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Productivity-Real Exchange Rate Nexus: Revisiting the Balassa-Samuelson Hypothesis for Emerging Asia

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ABSTRACT

This dissertation deals with the economics of equilibrium real exchange rates in the context of Balassa-Samuelson hypothesis. For developing economies of East and South Asia (ASEAN and SAARC), trend deviations in real exchange rate from long-run PPP are tested against their biased sectoral productivity patterns, under the theoretical predictions of Balassa-Samuelson model. The thesis undertakes a study of three inter-related dimensions of productivity-real exchange rate linkage: (a) identification of productivity as a key determinant of permanent deviations in real exchange rate from long-run PPP, (b) validity of hypothesis under alternative theoretical specifications, and (c) inclusion of demand-side shocks to see if this can buy any support for the proposed model.

The novelty of this study lies with the careful theoretical as well as empirical examination, conducted to verify the long-run association between sectoral productivity real exchange rate movements. To assess the robustness of results, distinction between traded and non-traded sectors of the real economy is made in a more definitive manner. Data inconsistencies across sectors as well as across countries in the form of uncommon data sources, inconsistent scheme of sectoral division and inadequately disaggregated sectors are addressed to ensure data reliability. Furthermore, two alternative schemes of sectoral classification are employed to examine the sensitivity of model estimates. Three alternative measures of real exchange rate are used, that dominantly comprise of nontradables prices, so that the internal mechanism of the Balassa-Samuelson model could be captured appropriately. Three alternative theoretical specifications of the Balassa-Samuelson model are tested, ranging from the most restrictive domestic version of the

model to the modified version, allowing for deviations in tradables prices from long-run PPP. The proposed model is also tested for its augmented version, by including a two demand-side determinants of real exchange rate, largely advocated in literature. Finally, various time-series and pooled data econometric techniques are applied for empirical verification, ranging from single equation to multivariate cointegration approaches, to test the consistency and robustness of estimates.

The results suggest that inter-country divergent sectoral productivity patterns do not exert any significant effect on the long-run real exchange rates for the ASEAN and SAARC countries in my sample, as predicted by Balassa (1964) and Samuelson (1964). The argument of the Balassa-Samuelson hypothesis that biased relative productivity of tradables at home will influence the overall price level of the country through nontraded sector prices and contribute to the long-run movements of real exchange rates does not hold valid. My findings are highly robust and successfully survived alternative sectoral classifications, different variants of real exchange rate measures, alternative theoretical specifications of the model and different econometric techniques. Empirical results for the standard (international) version of the hypothesis reveal that relative sectoral productivity differences across countries are inadequate in explaining the trend departures in the real exchange rates away from their long-run equilibrium. However, the Balassa-Samuelson hypothesis is more convincingly rejected when the domestic version of the model is tested. The results remain unchanged when the assumption of PPP, inherent to international Balassa-Samuelson model, is relaxed in favor of inter-country traded sector prices. This modified version of the model, allowing for trend deviations in international tradable prices from its PPP value, does not yield sizeable support, in favor of Balassa-Samuelson hypothesis for the sample countries. Furthermore, the empirical results reveal a clear departure of tradable prices from the long-run PPP, suggesting that this divergence is a potential reason for the non-existence of the Balassa-Samuelson effect. Finally, two demand-side factors, i.e., GDP per capita and government consumption spending, are augmented into Balassa-Samuelson model. However, their representation in the Balassa-Samuelson could gain only marginal support for the proposed productivity-real exchange rate relationship.

On the whole, I tend to reject the Balassa-Samuelson effect for emerging Asian countries due to inadequate empirical evidence in support of the hypothesis. Irrespective of alternative sectoral divisions, real exchange rate measures, model specifications or estimation procedures, sectoral productivity patterns are rarely found to cause significant real exchange rate appreciation in long-run for Asia.

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CHAPTER 1: INTRODUCTION

This dissertation deals with the economics of equilibrium real exchange rates. Previous studies have suggested that sustained departures of real exchange rates from its equilibrium values can occur due to real disturbances in the economy like productivity shocks, terms of trade shocks, oil shocks, government expenditure shocks, and so on. In my thesis, I focus on one such disturbance, the productivity shock. Unbalanced sectoral productivity growth across countries has been theoretically argued to be one of the explanations of the trend departure of real exchange rate from its long-run equilibrium. In the literature, this is known as the Balassa-Samuelson (BS) hypothesis. In this thesis, I empirically investigate the BS hypothesis for East and South Asian emerging economies.

My interest in East and South Asia stems from multiple sources. East and South Asia is widely recognized as being the growth driver of the world economy (Global Financial Stability Report, IMF, October, 2014). What happens there has consequences for the rest of the world. Further, the area has never really left the shadow of the Asian financial meltdown of 1997-1998. There is concern that this type of economic disaster could be repeated in the future. Perhaps fundamental changes in the economic structures underlying the region could minimize the chances of this event recurring. Finally, my personal background, and professional expertise, lies in this region, so I feel I am well familiar with the associated economic backstories.

Real exchange rate equilibrium is a longstanding notion which has always been imperative in explaining the mechanism of economic and financial stability of a region. The descriptions on long-run dynamics of exchange rate can be traced back to 1918 and 1922 when Gustav Cassel first introduced the concept of Purchasing Power Parity (PPP). The theory is regarded to be the best manifestation of potentially equilibrium real exchange rates. In a macroeconomic environment where exchange rates and commodity prices are perfectly flexible and cross-country trade is absolutely frictionless, the relative prices between countries tend to be stable or stationary over time. This implies that relative prices of a common basket of goods across countries are equalized when quoted in a common currency, assuming no market frictions and rigidities.

However, thinking of a well-connected, perfectly competitive and frictionless world economy is an idealist approach. For the same reason, real exchange rate can permanently deviate from its equilibrium. Financial products can be exchanged seamlessly in international markets but the same is not true for the real economy. A number of goods (particularly services) cannot be traded easily across borders. Thus, the compatibility between the general price levels in two countries with their bilateral exchange rate will reflect not only PPP-related forces but also the effects imparted by relative price differences of potentially more tradable and less tradable goods. Harrod (1933), Balassa (1964) and Samuelson (1964) extended the PPP proposition to account for inter-country productivity gaps as a potential source of causing long-run real exchange rate misalignments. In their classic articles *The Purchasing Power Parity Doctrine: A Reappraisal* by Bela Balassa and *Theoretical Notes on Trade Problems* by Paul A. Samuelson published in 1964, for the first time the authors formally argued why an absolute version of PPP is unlikely to hold in the long-run. They introduced the role of a cross-country sectoral productivity growth differential between traded and nontraded sectors as an important factor responsible for bringing systematic biases into the relationship between inter-country relative prices and equilibrium real exchange rates. When a country is more productive than the other, the former's nontraded sector tends to face a sharper rise in price level relative to the latter, resulting in a real appreciation of the former country's real exchange rate.

According to their Balassa and Samuelson's theory, a developed economy tends to be technologically more advanced compared to a less-developed economy. However, this technological advancement is not uniform across all sectors of the economy. It tends to be more pronounced in the traded sector of the economy relative to the nontraded one. Assuming purchasing power parity in the tradable sector with no transportation costs, trade restrictions, market frictions or rigidities, the prices of tradables from various countries will be equalized in international markets. However, this will not be true for nontradables prices which are determined my domestic market forces. Thus, if the home country is more productive in the traded sector relative to its trading partner, the domestic wage rate of the traded sector in that country will increase. The wage rise will bring cost-push inflation for the nontraded sector and will pull their prices upwards. This causes the domestic real exchange rate to appreciate against its trading partners. Hence, through shifts in internal (sectoral) price mechanism, the long-run exchange rate will deviate from the PPP.

Balassa and Samuelson's theoretical findings, that cross-country sectoral productivity growth bias is a powerful supply side explanation of exchange rate divergence from its PPP position in the long-run, led to verifiable empirical tests. However, in the empirical literature, the evidence on the valid existence of BS effect is highly mixed. Though the hypothesis is mostly tested for European and OECD economies, there is a vast contrast in findings, reported by different studies. The potential reasons for such conflicting results range from model specifications, econometric tests employed, data and variables used in the study, or testing the different versions of the hypothesis itself. In my thesis, I investigate at all these empirical issues.

My study contributes to the existing literature on the empirical testing of the BS hypothesis. My study focusses on the East and South Asian countries where there are only a handful of studies. I contribute to the literature by testing multiple versions of the BS hypothesis using a wide range of econometric techniques. A vast majority of studies investigating the BS hypothesis conduct country-by-country examination using different time-series econometric procedures (see Heston et al., 1994; Ito et al., 1999; Kakkar, 2000; Wang and Dunne, 2003; Lommatzsch and Tober, 2004; Mihaljek and Klau, 2004; Bogoev et al., 2008; Lothian and Taylor, 2008; Thomas and King, 2008; Chowdhury, 2012; Boreo et al., 2015). However, pooled data estimation methods have also been used in the recent literature (see Choudhri and Khan, 2005; Lee and Tang, 2007; Chong et al., 2012; Dumrongrittikul, 2012; Kakkar and Yan, 2012; Ricci et al., 2013; Wang et al., 2016). In this dissertation, I use both time-series and panel data cointegration techniques to test the robustness of the BS hypothesis. Each approach is further complemented by single equation residual based cointegration estimations and multivariate maximum likelihood based cointegration models. It is evident from the literature that the BS hypothesis provides conflicting results when analysed using alternative estimation techniques. However, to my knowledge, very few studies have verified the BS model under alternative econometric techniques. Thus, I employ a rigorous empirical framework in this dissertation to test the robustness and consistency of the BS estimates against alternative cointegration tests. Next, I briefly describe rest of the chapters of my dissertation.

In Chapter 2 of the dissertation, theoretical formulation of Balassa Samuelson model is discussed. This will help readers better understand those economic channels through which supply-side forces in the economy work. This conceptual building blocks of BS model may also prove to be useful while understanding different variants of the model and model variables empirically tested in the forthcoming chapters.

Chapter 3 discusses the data and the construction of the variables for my empirical exercise. I discuss in detail the formulation of the different measures of real exchange rate, sectoral prices and sectoral productivity. In this regard, special attention is paid to sectoral output and employment data conversions when distinguishing the economy into traded and nontraded sectors. The chapter explains how I handle the issues of inconsistencies in the data across sectors as well as countries and over time. The issue is intrinsic, but rarely discussed in the existing literature. I also discuss the issues regarding the division of economy into traded and nontraded sectors.

Chapter 4 will provide readers with a reasonably detailed insight into the existing empirical work done on the exchange rate behaviour in the context of the BS phenomenon. Since my work focusses on Asian economies, I will mainly focus my literature review on the Asian transition and emerging economies. I will critically evaluate the studies with reference to (a) different classification of sectors amongst tradables and nontradables, (b) different sectoral prices and productivity measures employed, (c) the econometric methodology employed to empirically verify the BS effect and (d) critical evaluation of the theoretical foundations of the model.

Chapter 5 empirically investigates whether sectoral productivity gap between tradables and nontradables can account for the permanent real exchange rate deviations in East and South Asia where U.S. is the reference country. The hypothesis is tested under the most basic and standard theoretical specification of the BS model. The analysis employs the most aggregated sectoral division prescribed in the literature classifying the real economy into four broad sectors. However, consistency and reliability of results is not compromised. Special care is taken to ensure robustness of the results. I use alternative measures of real exchange rate and various econometric methods to empirically test the BS hypothesis.

Chapter 6 primarily aims at examining the robustness of the results in chapter five. The model of Chapter 5 is tested using more disaggregated sectoral level data giving better coverage of industries under traded and nontraded sectors. This measure will serve as an additional robustness test on BS estimates. This will reveal how sensitive the results are to finer sectoral classifications.

Chapter 7 re-investigates the productivity-real exchange rate relationship by paying more attention to the internal transmission mechanism of the BS theory. This will reveal the relative significance of the domestic version of the model against its international counter-part.

Chapter 8 estimates the extended version of conventional BS model relaxing the assumption of PPP for tradables and allowing for lack of long-run co-movement between international tradables. This will not only provide an opportunity to obtain more factual estimates for the model, but will also give us an insight into the role of tradables prices in determining real exchange rate movements as emphasized by many earlier and recent empirical studies.

Chapter 9 includes some demand-side factors in the traditional BS model. Controlling for the inexistence of PPP between inter-country tradables creates room for building a more inclusive model of productivity-real exchange rate relationship by allowing for the additional role of demand-side real shocks. The primary objective of this chapter is to see if the inclusion of demand side determinants in BS model makes any difference to my prior findings on the invalid existence of BS effect for Asia. Selective demand-side determinants are augmented in the relaxed version of BS model (used in Chapter 8) and tested empirically as a plausible explanation of trend deviations of real exchange rate from its PPP-based long-run equilibrium.

Chapter 10 summarises my research findings and concludes that there is insufficient empirical evidence that BS hypothesis holds for East and South Asian countries. The findings are robust to different versions of real exchange rate, alternative schemes of sectoral division, alternative assumptions to the theoretical specifications of the model and different econometric estimation techniques.

CHAPTER 2: THEORETICAL EXPOSITION OF THE BALASSA-SAMUELSON HYPOTHESIS

The purpose of this chapter is to present the theoretical foundations of the Balassa-Samuelson (BS) model that will provide a benchmark for understanding the empirical analysis in the forthcoming chapters.

Consider a small open economy that produces two composite goods, tradables and nontradables. Output in the economy is produced by constant returns to scale (CRS) production function using two inputs, capital and labour. For simplicity, it is assumed that the production function is of Cobb-Douglas type:

(2.1-a) $Y^T = A^T L^{T\delta} K^{T(1-\delta)}$

(2.1-b) $Y^{NT} = A^{NT} L^{NT\gamma} K^{NT(1-\gamma)}$

where the superscript *T* denotes the traded sector and *NT* denotes the nontraded sector,

Y^T, Y^{NT} denote the outputs of traded and nontraded sectors respectively,

A^T, A^{NT} denote the technical efficiencies of traded and nontraded sectors respectively,

L^T, L^{NT} denote the labour supplies of traded and nontraded sectors respectively,

K^T, K^{NT} denote capital stocks of traded and nontraded sectors respectively,

δ, $(1 - \delta)$ are the production elasticities of the tradable sector's labour and capital respectively,

γ, $(1 - \gamma)$ are the production elasticities of the nontradable sector's labour and capital respectively, and

$0 < \gamma, \delta < 1$.

The model assumes that labour is mobile across sectors, but not across countries. However, capital is mobile across both sectors and countries. Factor markets are perfectly competitive. The total domestic labour supply is exogenously fixed at $L = L^T + L^{NT}$. Labour mobility across sectors ensures that workers earn the same wage rate W in either sector. There is no economy-wide resource constraint for capital. Thus, the capital's domestic rate of return is equal to the world interest rate R.

2.1 Domestic Version of the Balassa-Samuelson Hypothesis

The representative firm is a profit maximiser. Let P be the relative price of nontradables to tradables, i.e., $\left(\frac{P^{NT}}{P^T}\right)$. Assuming (for simplicity) a constant world interest rate, the profit functions in the tradable and nontradable sectors, measured in units of tradables, are given as:

(2.2-a) $\pi^T = Y^T - W^T L^T - RK^T$

(2.2-b) $\pi^{NT} = PY^{NT} - W^{NT} L^{NT} - RK^{NT}$

Profit maximization yields the following first-order conditions from equations (2.2-a) and (2.2-b).

For the traded sector:

(2.3-a) $\frac{\partial \pi^T}{\partial L^T} = 0 => \delta A^T L^{T(\delta-1)} K^{T(1-\delta)} = W^T$

(2.3-b) $\frac{\partial \pi^T}{\partial K^T} = 0 => (1-\delta)A^T L^{T\delta} K^{T(-\delta)} = R$

Similarly, for the nontraded sector:

(2.3-c) $\frac{\partial \pi^{NT}}{\partial L^{NT}} = 0 => \gamma P A^{NT} L^{NT(\gamma-1)} K^{NT(1-\gamma)} = W^{NT}$

(2.3-d) $\frac{\partial \pi^{NT}}{\partial K^{NT}} = 0 => (1-\gamma) P A^{NT} L^{NT\gamma} K^{NT(-\gamma)} = R$

Equations (2.3-a) – (2.3-d) imply the following:

(2.4-a) $\mathrm{MPL}^{\mathrm{T}} = \delta A^T \left(\frac{K^T}{L^T}\right)^{1-\delta} = W^T$

(2.4-b) $\mathrm{MPK}^{\mathrm{T}} = (1-\delta)A^T \left(\frac{K^T}{L^T}\right)^{-\delta} = R$

(2.4-c) $\mathrm{MPL}^{\mathrm{NT}} = \gamma P A^{NT} \left(\frac{K^{NT}}{L^{NT}}\right)^{1-\gamma} = W^{NT}$

(2.4-d) $\mathrm{MPK}^{\mathrm{NT}} = (1-\gamma) P A^{NT} \left(\frac{K^{NT}}{L^{NT}}\right)^{-\gamma} = R$

Re-arranging equations (2.4-b) and (2.4-d) yields:

(2.5-a) $\left(\frac{K^T}{L^T}\right) = \left(\frac{R}{(1-\delta)A^T}\right)^{-(\frac{1}{\delta})}$

(2.5-b) $\left(\frac{K^{NT}}{L^{NT}}\right) = \left(\frac{R}{(1-\gamma)PA^{NT}}\right)^{-(\frac{1}{\gamma})}$

Substituting equations (2.5-a) and (2.5-b) into equations (2.4-a) and (2.4-c) respectively, and re-writing in terms of sectoral wages yield:

(2.6-a) $W^T = \delta A^T \left(\frac{R}{(1-\delta)A^T}\right)^{-(\frac{1-\delta}{\delta})}$

(2.6-b) $W^{NT} = \gamma P A^{NT} \left(\frac{R}{(1-\gamma)PA^{NT}}\right)^{-(\frac{1-\gamma}{\gamma})}$

As wages are assumed to be equal across sectors, i.e., $W^T = W^{NT}$, equations (2.6-a) and (2.6-b) yield:

(2.7) $\delta A^T \left(\frac{R}{(1-\delta)A^T}\right)^{-(\frac{1-\delta}{\delta})} = \gamma P A^{NT} \left(\frac{R}{(1-\gamma)PA^{NT}}\right)^{-(\frac{1-\gamma}{\gamma})}$

Taking logarithm on both sides of equation (2.7) and re-arranging terms:

(2.8) $\left(\frac{1}{\gamma}\right) \ln P = \left(\frac{1}{\delta}\right) \ln A^T - \left(\frac{1}{\gamma}\right) \ln A^{NT} + C$

where $C = \left(\frac{(1-\gamma)}{\gamma} - \frac{(1-\delta)}{\delta}\right) \ln R + \ln \delta + \frac{(1-\delta)}{\delta} \ln(1-\delta) - \ln \gamma - \frac{(1-\gamma)}{\gamma} \ln(1-\gamma)$

Multiplying both sides of equation (2.8) by γ and using the definition of $\ln P$ yields:

(2.9) $p = p^{NT} - p^T = \left(\frac{\gamma}{\delta}\right) a^T - a^{NT} + c$

where small letters represent the variables in their logarithmic form and the constant $c = \gamma C$.

Equation (2.9) represents the "*domestic version*" of the BS model and explains how the relative movement of sectoral prices are driven by the relative movements of technical efficiencies (also referred to as productivities) in each sector within a country. If the non-traded sector is equally or relatively more labour-intensive than the traded sector, i.e. $\gamma/_\delta \geq 1$, faster productivity growth in traded sector compared to the non-traded sector will cause the non-tradable sectoral prices to rise.

2.2 International (Standard) Version of the Balassa-Samuelson Hypothesis

To illustrate the international version of the BS hypothesis, assume the following about prices in two countries: traded goods have the same price at home and abroad and nontraded goods have distinct prices at home and abroad. In other words, *Purchasing Power Parity* (*PPP*) holds for the traded sector, but not for the nontraded sector. It is further assumed that the domestic and foreign price levels are geometric averages of the tradable and nontradables. Thus in logarithmic form:

(2.10-a) $p = (1-\beta)p^{T} + \beta p^{NT}$

(2.10-b) $p^{*} = (1-\beta)p^{T*} + \beta p^{NT*}$

The small letters represent the variables in logarithmic form. The real exchange rate (RER) is defined as a product of nominal exchange rate (E) between two countries and the relative prices of foreign ($P*$) and home (P) countries. Thus in logarithmic form:

(2.11) $rer = e + p^{*} - p$

Assumption of PPP in the tradable sector implies:

(2.12) $p^{T} = e + p^{T*}$

Using equation (2.10-a), (2.10-b), (2.11) and (2.12), the real exchange rate can be re-written in logarithmic form as:

(2.13) $rer = (e + p^{T*} - p^{T}) - \beta\{(p^{NT} - p^{T}) - (p^{NT*} - p^{T*})\}$

Substituting equation (2.9) in (2.13), the real exchange rate can now be written as:

(2.14-a) $rer = (e + p^{T*} - p^{T}) - \beta \left\{ \left({}^{\gamma}/_{\delta}\, a^{T} - a^{NT} \right) - \left({}^{\gamma}/_{\delta}\, a^{T*} - a^{NT*} \right) \right\}$

If Purchasing Power Parity (PPP) holds in the traded sector, the term $(e + p^{T*} - p^{T}) = 0$. Thus equation (2.14-a) can be written as:

(2.14-b) $rer = -\beta \left\{ \left({}^{\gamma}/_{\delta}\, a^{T} - a^{NT} \right) - \left({}^{\gamma}/_{\delta}\, a^{T*} - a^{NT*} \right) \right\}$

Equation (2.14-b) represents the *Balassa-Samuelson* effect. Higher productivity growth in the traded sector than the non-traded one in the home country causes an appreciation of the home country real exchange rate, i.e. a decreasing real value of home currency relative to the foreign one. In the later chapters, we test this proposition more formally, using econometric models.

CHAPTER 3: DATA DISSEMINATION AND SECTORAL CLASSIFICATION

In Chapter One, I explain the importance of Balassa-Samuelson (BS) hypothesis in understanding real exchange rate movements in the long-run. In Chapter Two, I present a theoretical framework for understanding the BS hypothesis and provide testable equations for the empirical analyses in the forthcoming chapters. This chapter will take the readers one step ahead. It will provide all the necessary information on data that are needed to verify the validity of the BS hypothesis. In particular, I will discuss the challenges associated with the measurement of the variables and discuss how I address those issues. I pay special attention to two things: (i) data inconsistencies that arise over time and across countries, and (ii) division of real economy into traded and nontraded sectors, both of which have often been largely overlooked in the empirical studies investigating the BS hypothesis.

3.1 Measuring the Regressand and the Regressor: Real Exchange Rate and its Variants

The real exchange rate between two countries is defined as the product of their bilateral nominal exchange rate and relative prices.

(3.1.1-a) $$RER_t = \frac{E_t P_t^*}{P_t}$$

A logarithmic transformation of the variables yield:

(3.1.1-b) $$rer_t = e_t + p_t^* - p_t$$

where

$rer_t =$ logarithm of bilateral real exchange rate between home and foreign economies at time *t*, defined as units of home goods per unit of foreign good at time *t*.

$e_t =$ logarithm of bilateral nominal exchange rate between home and foreign economies at time *t*, defined as units of home currency per unit of foreign currency at time *t*.

$p_t^* =$ logarithm of aggregate price level of foreign economy at time *t* expressed in terms of foreign currency. It represents the expenditures made on the purchase of a standardized consumption basket at time *t*. The consumption basket largely comprises of nontradable goods and services.

$p_t =$ logarithm of aggregate price level of home economy at time *t* expressed in terms of home currency. It is characterized by the same compositional features as stated for p_t^*.

An increase in the value of rer_t will be interpreted as real exchange rate depreciation of home country vis-à-vis the foreign country at time *t*.

3.1.1 Measures of Real Exchange Rate

Throughout this dissertation, the empirical verification of the BS hypothesis is conducted by employing three different measures of real exchange rate. The three measures are different from each other on the basis of price measures used to deflate exchange rate. Selection of relevant price measure is a vital issue as it bears numerous direct and indirect effects on the final estimates of the Balassa-Samuelsson effect. Ideally, the price measure used should be comparable to the relevant basket of goods and services across countries. Moreover, it should also reflect the long-run permanent price trend instead of short-term temporary fluctuations. Nevertheless, no available price measure fits this criteria and there is no consensus on the price measure that should be used in constructing the real exchange rate.

My goal is to select a price measure that may serve as the best approximation of nontradable goods prices of a country. Since no price measure reflects only components of nontradable goods prices, I employ, for robustness, three alternative measures of prices that have a large representation of nontraded sector goods and services in the price index.

(i) CPI Deflated Real Exchange Rate:

CPI is one of the most widely used indices for measuring RER (see Taylor, 1996; Egert, 2002; Mihaljek & Klau, 2008; Dumrongrittikul, 2012). CPI represents the overall cost (weighted average) of a fixed basket of goods and services bought by a typical consumer relative to the price of the same basket in a base year. By including a broad range of goods and services (dominantly nontradables like housing, consumer services, public administration, defence, medical care, etc.) in the fixed basket, CPI can obtain an accurate estimate of the cost of living. It is important to remember that the CPI is not a dollar value like GDP; it is an index number which captures the overall change in the price level from the base year. A change in this index over time should reflect the position of nontraded sectors over time. However, CPI contains some traded goods in its basket. If the proportion of traded goods in the basket is high, it might not be the ideal indicator for comparison.

The CPI deflated RER is defined as:

(3.1.2) $rer\_cpi_t = e_t + p^*\_cpi_t - p\_cpi_t$

where

$* = \text{U.S.},$

rer_cpi_t = logarithm of CPI based real exchange rate at time *t*,

e_t = logarithm of bilateral nominal exchange rate between home and U.S. at time *t*, defined as units of home currency per unit of USD at time *t*,

$p^*\_cpi_t$ = logarithm of CPI of U.S. at time *t*. CPI uses year 2005 as the reference year,

p_cpi_t = logarithm of CPI of home at time *t*. CPI uses year 2005 as the reference year.

For constructing CPI based real exchange rate, I take both series (at annual frequency) from International Financial Statistics (IFS) database published by International Monetary Fund (IMF). The latest version of IFS provides CPIs with year 2010 as the reference year. I rebase the series to year 2005 to make the CPI based real exchange rate series comparable to the other variables in my regressions.

(ii) GDP Deflator-Based Real Exchange Rate:

My second measure of RER involves GDP deflators as a proxy for price indices at home and foreign. The GDP deflator is considered superior to CPI as a price measure in this case since it covers a broader range of commodities and services that are nontraded. In addition, the GDP deflator is highly proficient in capturing the effects of productivity shocks on real exchange rate, provided the component of regulated prices of two type of industries that constitute lion share of consumption expenditures (food and services in particular) are controlled for (Jazbec, 2002; Egert et al., 2003).

The GDP deflator based RER is defined as:

(3.1.3) $$rer\_def_t = e_t + p^*\_def_t - p\_def_t$$

where

$* = \mathrm{U.S.}$,

$rer\_def_t =$ logarithm of GDP deflator based real exchange rate at time *t*,

$e_t =$ logarithm of bilateral nominal exchange rate between home and U.S. at time *t*, defined as units of home currency per unit of USD at time *t*,

$p^*\_def_t =$ logarithm of GDP deflator of U.S. at time *t*,

$p\_def_t =$ logarithm of GDP deflator of the home at time *t*.

The reference year for the indices is 2005.

The GDP deflator series are constructed by using annual country-level GDP data at current and constant market prices as given below.

(3.1.4-a) $$p^*\_def_t = ln\left[\left(\frac{NGDP_t^*}{RGDP_t^*}\right) * 100\right]$$

(3.1.4-b) $$p\_def_t = ln\left[\left(\frac{NGDP_t}{RGDP_t}\right) * 100\right]$$

where

$NGDP_t =$ Nominal GDP (in million national currency) at time *t*,

$RGDP_t =$ Real GDP (in million national currency) at time *t*.

GDP series are sourced from the national accounts main aggregates database of United Nations Conference on Trade and Development (UNCTAD).

(iii) Non-Tradables Prices Based Real Exchange Rate:[1]

The third measure of real exchange rate is based on an index of nontradables prices. The real exchange rate is defined as:

$$(3.1.5) \qquad rer\_def\_nt_t = e_t + p^*\_def\_nt_t - p\_def\_nt_t$$

where

$* = \text{U.S.}$,

$rer\_def\_nt_t$ = logarithm of nontradables prices based real exchange rate at time *t*,

e_t = logarithm of bilateral nominal exchange rate between home and U.S. at time *t*, defined as units of home currency per unit of USD at time *t*,

$p^*\_def\_nt_t$ = logarithm of nontradables price deflator of U.S. at time *t*,

$p\_def\_nt_t$ = logarithm of nontradables price deflator of the home country at time *t*.

The base year for the indices is 2005.

Constructing an index of real exchange rate based on nontradables price deflator requires data on sectoral prices. The time series involved in the construction of sectoral price deflators are sourced from national accounts main aggregates database of United Nations Conference on Trade and Development (UNCTAD). The database provides annual data on nominal and real output of

[1] UNCTAD provides nominal and real GDPs by splitting the economy into seven sectors. For Chapter 6, the measure is constructed by identifying the nontradable sectors from all seven sectors (see Section N for the scheme of division between tradables and nontradables). However, for Chapter 5, the seven sectors are aggregated into four broad sectors, i.e., agriculture, manufacturing, industry and services. Agriculture and manufacturing are grouped together to represent tradable sector of each country whereas industry and services together represent nontradable sectors of the real economy.

the economy by splitting it into seven distinct sectors (Section 3.2 further elaborates the sectoral data and the scheme of sectoral division used in this study).

The price series involved in the construction of this version of real exchange rate are weighted averages of prices of nontradable sector value-added, where the weights are the corresponding industry shares in total value added of the country.

(3.1.6-a) $$p^{*}\_def\_nt_{t} = ln\left[\sum_{s=1}^{k} w_{s} \left(\frac{NVA_{s,t}}{RVA_{s,t}}\right) \times 100\right]$$

(3.1.6-b) $$p\_def\_nt_{t} = ln\left[\sum_{s*=1}^{k} w_{s*}^{*} \left(\frac{NVA_{s*,t}^{*}}{RVA_{s*,t}^{*}}\right) \times 100\right]$$

where

$* = \mathrm{U.S.}$,

$NVA =$ Nominal Value Added (in million national currency) at time *t*,

$RVA =$ Real Value Added (in million national currency) at time *t*,

$w_{s} =$ weight by share of each nontradaded industry in home total value added,

$w_{s*}^{*} =$ weight by share of each nontradaded industry in U.S. total value added,

$s = 1,2, \ldots \ldots, k$ nontradable industries at home,

$s^{*} = 1,2, \ldots \ldots, k$ nontradable industries at U.S.

Similar to GDP series, the value added data series are sourced from the national accounts main aggregates database (UNCTAD)

In the construction of price and productivity series, I allow the sectoral weights to vary across years as well as across countries, as the composition of sectoral output for the countries in my sample have shifted dramatically during the course of sample study period. Consequently, this is reflected in the sectoral contribution to national output which also has evidently changed during this time. Thus, in the construction of prices and productivity data series, the sectoral weights need to match the growing share of nontradables in domestic GDP (if there is any), responsible for giving an upward push to a country's overall inflation, and thus, may help in capturing the plausible existence of BS effect more proficiently. The role of growing relative share of nontradables in the consumption basket in explaining the BS mechanism has been much emphasized in studies on Europe (Egert et al., 2003; and Mihaljek and Klau, 2004).

3.1.2 Measures of Relative Prices

I extend the empirical testing of the simple BS hypothesis by considering two different extensions of the hypothesis:

(a) Whether the BS hypothesis holds within the domestic economy, i.e., testing the "*domestic version*" of the BS hypothesis, and

(b) Is there any empirical support for the BS hypothesis when the fundamental assumption of Purchasing Power Parity (PPP) between inter-country tradable prices is relaxed. This assumption is empirically challenged by many authors (Isard, 1977; Engel, 1995; Canzoneri et al., 1999).

Testing the two extensions of the BS hypothesis require me to obtain "*intra-country*" and "*inter-country*" relative sectoral price measures. These relative price measures are different from the three variants of real exchange rate discussed in the preceding sections. Below, I briefly introduce the basic structure of the modified versions of the BS hypothesis along with explanations on respective price measures. The empirical estimates for the two models are discussed in details in Chapters 7, 8 and 9 later.

(i) Price Differential between Intra-Country Non-Traded and Traded Sectors:

A considerable number of studies have empirically tested the "*domestic version*" of the BS hypothesis (see Egert, 2002; Lee and Tang, 2003; Mihaljek and Klau, 2004; Thomas and King, 2008). These studies state that the relationship between domestic relative productivities and relative sectoral prices play a key role in driving the BS mechanism. In the literature, the domestic version of BS hypothesis is known as the "*Baumol-Bowen effect*". Baumol and Bowen (1966) observed the relative prices of service-intensive goods rising over time. On the other hand, the productivity growth of such industries tends to be slower than that in more capital-intensive manufacturing industries.

The empirical verification of the domestic version of the BS model requires relative domestic sectoral prices as regressand in the model. For sectoral prices, the only reliable and consistent data series available are sectoral VA deflators. For generating internal price ratio, I used two aggregate sectoral deflators for nontradable and tradable sectors, respectively. The internal price differential is expressed as:

(3.1.7) $$p\_int_t = p_t^{NT} - p_t^T$$

where

p_int_t = internal price diffential at time *t*

p_t^{NT} = logarithm of nontradable prices of home at time *t*

p_t^T = logarithm of tradable prices of home at time *t*

The sectoral price deflators are calculated as:

(3.1.8-a) $$p_t^{NT} = ln\left[\sum_{NT=1}^{k} w_{NT} \cdot \left(\frac{NVA_{NT,t}}{RVA_{NT,t}}\right) \times 100\right]$$

(3.1.8-b) $$p_t^T = ln\left[\sum_{T=1}^{k} w_T \cdot \left(\frac{NVA_{T,t}}{RVA_{T,t}}\right) \times 100\right]$$

where

$w_{NT} =$ weight by share of each nontradable industry in total VA of home

$w_T =$ weight by share of each tradable industry in total VA of home

$NVA =$ Nominal Value Added (in million national currency) at time *t*

$RVA =$ Real Value Added (in million national currency) at time *t*

Sectoral current and constant VA data are at annual frequency and I follow similar construction guidelines as reported in Section 3.1.1 (iii). The data series are obtained from national accounts main aggregates database (UNCTAD) that I use in the construction of the price series in Section 3.1.1 (iii).

(ii) Price Differential between Inter-Country Non-Traded and Traded Sectors

One of the vital assumptions underlying the BS hypothesis is that Purchasing Power Parity (PPP) holds between the traded sectors of two countries. In other words, the BS effect assumes that domestic and foreign tradables are perfect substitutes. If PPP in the tradable sector holds, a productivity improvement in the tradable sector at home will drive up the wages for the entire economy as it also increases the labour cost for the nontradable sector. This causes the relative price of the nontradable sector at home to rise vis-à-vis the foreign country leading to an appreciation of the real exchange rate in the home country. However, if PPP does not hold in the tradable sector, imbalanced productivity growth amongst sectors can affect real exchange rate through both inter-country relative price of the nontradables as well as the tradable sector (please refer to Chapter 8 for detailed discussions on this subject).

Testing this version of the BS hypothesis requires two types of inter-country relative prices in the model. The first is the inter-country relative prices of nontradables as the regressand and the second is the inter-country relative prices of tradables as a regressor. These are defined as:

(3.1.9-a) $$p_t^{NT*} - p_t^{NT} = e_t + va\_def\_nt_t^* - va\_def\_nt_t \equiv rer\_def\_nt$$

(3.1.9-b) $$p_t^{T*} - p_t^{T} = e_t + va\_def\_t_t^* - va\_def\_t_t \equiv rer\_def\_t$$

where

$e_t =$ bilateral nominal exchange rate between home and U.S. at time *t*

$p_t^{NT*} - p_t^{NT} =$ price differential between U.S. and home country in the nontraded sector at time *t*

$p_t^{T*} - p_t^{T} =$ price differential between U.S. and home country in traded sector at time *t*

$$va\_def\_nt_t^* = ln\left[\sum_{NT*=1}^{k} w_{NT*}^* \cdot \left(\frac{NVA_{NT*,t}^*}{RVA_{NT*,t}^*}\right) \times 100\right]$$

$$va\_def\_nt_t = ln\left[\sum_{NT=1}^{k} w_{NT} \cdot \left(\frac{NVA_{NT,t}}{RVA_{NT,t}}\right) \times 100\right]$$

$$va\_def\_t_t^* = ln\left[\sum_{T*=1}^{k} w_{T*} \cdot \left(\frac{NVA_{T*,t}^*}{RVA_{T*,t}^*}\right) \times 100\right]$$

$$va\_def\_t_t = ln\left[\sum_{T=1}^{k} w_{T} \cdot \left(\frac{NVA_{T,t}}{RVA_{T,t}}\right) \times 100\right]$$

where $* = \text{U.S.}$

$w_T =$ weight by share of each tradable industry in total VA of home

$w_{NT} =$ weight by share of each nontradable industry in total VA of home

w_{T*} = weight by share of each tradable industry in total VA of U. S.

w_{NT*} = weight by share of each nontradable industry in total VA of U. S.

Since, the two price measures are constructed from nontradable and tradable sector price deflators, the data source and the specifications for current and constant sectoral VAs are the same as reported in Section 3.2.1 (iii).

3.1.3 Measuring the Regressor: Productivity Gap and its Variants

The core explanatory variable of the BS mechanism is the sectoral productivity differential across sectors and between countries. Thus, its precise and accurate measurement is extremely important for the empirical analysis. In the earlier studies on the BS hypothesis, a vast majority of authors use Average Productivity of Labour (APL) as a proxy for productivity (see Heston et al., 1994; Chinn, 2000; Mihaljek & Klau, 2004; Lee & Tang, 2007; Thomas and King, 2008); Dumrongrittikul, 2012; Ricci et al.,2013). The popularity of this measure amongst researchers can be attributed to its computational ease and consistency in measurement across different countries.

However, Total Factor Productivity (TFP) has also been considered by some authors as a measure of productivity (see Asea, 1994; De Gregorio et al., 1994; Kakkar, 2002; Lee & Tang, 2007; Olson, 2009; Kakkar and Yan, 2012). TFP is exogenous to investment dynamics and capital accumulation whereas APL is endogenous to investment dynamics. However, the cross-country sectoral productivity differentials calculated by using APL also take into account substantial proportions of TFP and investment dynamics differential. APL is also capable of capturing the difference in capital costs more efficiently (see Lee and Tang (2007)). As a result, APL is generally preferred over TFP when it comes to measuring cross-country sectoral productivity gaps. In my thesis, APL is chosen as the proxy measure of productivity in this analysis. Unavailability of sectoral level capital formation data for Asian countries prevented me from employing TFP for my analysis.

The forthcoming parts of this section distinguish between two variants of sectoral productivity gap at inter-country and intra-country levels and are supplemented by detailed discussions on their nature, construction and sources.

(i) Productivity Gap between Inter-Country Traded and Non-Traded Sectors[2]

The BS hypothesis establishes a long-run relationship between inter-country sectoral productivity differences and RER movement. This is often referred to as the "*international version*" of the model (see Chapter 5 for full specification of the model). I use this term to distinguish it from the "*domestic version*" of the model which I test in my later chapters (Chapter Seven).

Corresponding to RER and inter-country relative price measures as regressand (discussed in preceding sections), the international version of BS model requires a measure of sectoral productivity differentials between countries. I proxy this measure with the average productivity of labour in traded and nontraded sectors. Average Productivity of Labour (APL) is the amount of real value added produced by each unit of employed labour. The productivity measure in its final form can be written as the inter-country difference between relative productivity of labour in traded and nontraded sectors. Mathematically, this is written as:

$$\tilde{a}_t = -[(a_t^T - a_t^{NT}) - (a_t^{T*} - a_t^{NT*})] \tag{3.2.1}$$

The sectoral APLs are written as:

[2] Please refer to footnote 1 for explanation on the other version of the series.

(3.2.2-a) $$a_t^T = APL_t^T = \sum_{T=1}^{k} w_T \cdot \left[\left(\frac{RVA_{T,t}}{Employment_{T,t}}\right)\right]$$

(3.2.2-b) $$a_t^{NT} = APL_t^{NT} = \sum_{NT=1}^{k} w_{NT} \cdot \left[\left(\frac{RVA_{NT,t}}{Employment_{NT,t}}\right)\right]$$

(3.2.3-c) $$a_t^{T*} = APL_t^{T*} = \sum_{T*=1}^{k} w_{T*} \cdot \left[\left(\frac{RVA_{T*,t}}{Employment_{T*,t}}\right)\right]$$

(3.2.4-d) $$a_t^{NT*} = APL_t^{NT*} = \sum_{NT*=1}^{k} w_{NT*} \cdot \left[\left(\frac{RVA_{NT*,t}}{Employment_{NT*,t}}\right)\right]$$

where

$* = \text{U.S.}$

$\tilde{a}_t$ = relative productivity differential across traded and nontraded sectors and between home and U.S. at time *t*.

$(a_t^T - a_t^{NT})$ = logarithm of productivity difference between traded and nontraded sectors at home at time *t*.

$(a_t^{T*} - a_t^{NT*})$ = logarithm of productivity difference between traded and nontraded sectors in U.S. at time *t*.

RVA = Real Value Added at time *t*

$Employment$ = Employed labour force in millions

w_T = weight by share of each tradable industry in total VA

w_{NT} = weight by share of each nontradable industry in total VA

w_{T*} = weight by share of each tradable industry in total VA of U. S.

w_{NT*} = weight by share of each nontradable industry in total VA of U. S.

The annual sectoral real value added data is sourced from national accounts main aggregates database, UNCTAD. The sectoral employment (at annual frequency) is obtained through two distinct databases of International Labor Organization (ILO) published by ILO Department of Statistics (please refer to Section 3.2 for further details).

(ii) Productivity Gap between Intra-Country Traded and Non-Traded Sectors

In Section 3.1.2, I have briefly explained the theoretical rationale behind estimating the domestic version of BS hypothesis. The regressand in the model is the price ratio between domestic traded and nontraded sectors. The corresponding regressor of the model is the internal relative productivity of traded and nontraded sectors. This can be given as:

(3.2.5) Internal Productivity Gap $= a = a_t^T - a_t^{NT}$

where

a_t^T = logarithm of APL based tradables productivity of home at time *t*.

a_t^{NT} = logarithm of APL based nontradables productivity of home at time *t*.

TABLE 3.1: Data Description and Sources

Measure	Time-Series	Description	Source
Components of Real Exchange Rate and Relative Sectoral Prices			
Bilateral nominal exchange rate	Official exchange rate (LCU per U.S.$, period average)	Official exchange rate refers to the exchange rate determined by national authorities or to the rate determined in the legally sanctioned exchange market. It is calculated as an annual average based on monthly averages (local currency units relative to the U.S. dollar).	International Financial Statistics
Consumer Price Index (CPI) RER price deflator	Consumer Price Index (2005=100)	CPI represents changes in cost to the average consumer of acquiring a basket of goods and services that may be fixed or changed at specified intervals, such as yearly.	International Financial Statistics
GDP Deflator Real exchange rate price deflator	GDP Deflator (2005=100)	For each economy, the series of GDP deflator has been constructed with the help of below discussed nominal and real GDP series: *Nominal GDP*: U.S. Dollars at current prices and current exchange rates in millions *Real GDP*: U.S. Dollars at constant prices (2005) and constant exchange rates (2005) in millions. **Note**: GDP Deflator (nontradables) and internal relative price series are constructed using sectoral price deflators. The sectoral	National accounts main aggregates database (UNCTAD)

Measure	Time-Series	Description	Source
		price deflators follow same procedure for their construction, the one used for constructing aggregate value added deflators	
Components of Sectoral Productivity			
Sectoral Output (T and NT)	Average Labour Productivity (APL) *Sectoral Value Added (Mn) (Disaggregation into Seven Sectors)*	The sectoral productivity is measured in the form of Average Productivity of Labour (APL) which is a ratio of sectoral value added to the total labour employment in that sector. Thus, the construction of APL for tradables and nontradables involves the two time series; sectoral value added and sectoral labour employment. The description and the sources for two series are given below: *Agriculture*: Agriculture includes agriculture, hunting, forestry and fishing (it corresponds to ISIC Rev.3 divisions 01-05). *Mining & Utilities*: Corresponds to ISIC Rev 3 divisions 10-14 and 38-41. *Manufacturing*: Corresponds to ISIC Revision 3 divisions 15-37. *Construction*: Corresponds to ISIC Revision 3 division 45. *Whole sale & retail trade*: Corresponds to ISIC Revision 3 divisions 50-55.	National accounts main aggregates database (UNCTAD)

Measure	Time-Series	Description	Source
Sectoral Output (T and NT)	*Sectoral Value Added (Mn)* *(Disaggregation into Seven Sectors)*	*Transport, storage and communication*: It corresponds to ISIC Revision 3 divisions 60-64. *Other Activities*: It corresponds to ISIC Revision 3 div. 65-95. ________________________ *Agriculture*: Agriculture includes agriculture, hunting, forestry and fishing (it corresponds to ISIC Rev.3 divisions 01-05). *Manufacturing*: It corresponds to ISIC Rev.3 divisions 15-37. *Industry*: Includes mining and quarrying, electricity, gas and water supply, and construction. It corresponds to ISIC Rev.3 divisions 10-14, 40, 41 and 45. *Services*: Includes all other economic activities (it corresponds to ISIC Rev.3 divisions 50-55 and 60-99). ________________________	National accounts main aggregates database (UNCTAD)
Sectoral Employment (T and NT)	*Sectoral Employed Labour Force (Mn)* *(Disaggregation into Four Sectors)*	For each country, sectoral employment is measured by industry employment by kind of activity, i.e., employed labour force in each sector (million). For each subject economy, sectoral employment data follows ISIC-Rev.2, ISIC-Rev.3 and for U.S. the data is provided under NAICS-2007 sectoral classification scheme (from 2002-2013). Please refer to TABLE 3.6 to 3.8 to see detailed sectoral divisions under three sectoral classification schemes and the corresponding data period.	International Labour Organization (ILO). From year 1970 to 2008 the data is taken from LABORSTA (ILO) and from the year 2009 to 2013 the data is sourced from ILOSTAT (ILO)

TABLE 3.2: Summary Table for Real Exchange Rate, Price Differentials and Productivity Gap Measures

Measure	Mathematical Representation	Role in BS Model	Version of BS Model
CPI-based Real Exchange Rate	rer_cpi	Regressand	International-Standard
GDP Deflator based Real Exchange Rate	rer_def	Regressand	International-Standard
Non-Tradable Prices-based Real Exchange Rate	rer_def_nt	Regressand	International-Standard & Modified
Tradable Prices-based Real Exchange Rate	rer_def_t	Regressor	International-Standard & Modified
Inter-Country Productivity Gap	$\tilde{a}_t = -[(a_t^T - a_t^{NT}) - (a_t^{T*} - a_t^{NT*})]$	Regressor	International-Standard & Modified
Internal Price Differential	$p^{NT} - p^T$	Regressand	Domestic
Internal Productivity Gap	$a_t^T - a_t^{NT}$	Regressor	Domestic

3.2 Scheme of Differentiation between Traded and Non-Traded Sectors

Classifying an economy into two distinct sectors is a complex task. It is difficult due to the inconsistent schemes of sectoral division adopted by data collection authorities and misrepresentation or under-representation of industries that lead to potential problems in distinguishing between traded and nontraded sectors of the economy. The issue becomes more intricate when one finds disagreement between researchers on the correct method of sectoral division. These concerns are often overlooked in the existing literature. However, a vigilant treatment could make substantial differences in the final result of the BS hypothesis.

In the forthcoming sections, the first one talks about the approach adopted in this thesis to classify an economy into traded and nontraded sectors. The later section discusses the issues I face in the course of sectoral classification and the essential transformations I make to ensure comparability of output and employment data series.

3.2.1 Data Sources and Scheme of Sectoral Disaggregation

In the context of sectoral division, the discussion involves the two key variables – real output and employment. The two elements together generate the average productivity of labour at the sectoral level.

(i) Sectoral Division of Output:

Sectoral output data is sourced from national accounts main aggregates database, UNCTAD (please refer to Section 3.1 for details). The real economy is disaggregated into seven distinct sectors, which correspond to International Standard Industry Classification (ISIC) Revision – 3. This allows consistency of real output data across sectors as well as across countries.

BS hypothesis is empirically tested by disaggregating the economy into (a) seven sectors – agriculture, manufacturing, mining and utilities, construction, wholesale, retail trade and hospitality, transport and communication and other activities (please refer to Chapter 6 for empirical details) and (b) the seven sectors aggregated into four broad sectors – agriculture, manufacturing, industry and services (please refer to Chapter 5 for empirical details). TABLE 3.4 reports the scheme for disaggregating the economy into four and seven sectors. From the table, it can be seen that both types of sectoral divisions consistently follow the classification of ISIC Revision – 3. This ensures no abrupt shift in any sector.

The empirical analysis conducted in all the chapters (except chapter 5) involve data disaggregated into seven sectors. A number of prominent studies have also used less dis-aggregated sectors for empirically estimating the BS effect. These studies divide the economy into two or three broad sectors and recognize them as tradables or nontradables (see Simon and Kovacs, 1998; Halpern and Wyplosz, 2001; Fischer, 2004). Following their practice, I also test the BS model with less disaggregated sectoral division in Chapter 5 (see TABLE 3.4). I perform this analysis for two reasons: first, to test the consistency of my estimates against these studies and second, to test if a less disaggregated sectoral classification would generate a different result for the set of countries I analyse. However, studying BS hypothesis with more disaggregated data is more reliable and is found to produce substantially different results (see Alberola-Ila and Tyrväinen, 1998; Egert, 2003; Mihaljek and Klau, 2004).

To ensure consistency in sectoral output data across countries, I had to undertake some data transformations. From 1970 to 2013, UNCTAD publishes sectoral value added data under altering scheme of sectoral classifications changing from ISIC Revision 2 to ISIC Revision 3. Furthermore, the varying scheme of sectoral division is not consistent across countries. TABLE 3.3 displays the scheme of sectoral disaggregation of value added changing over time. Except for

Hong Kong and Pakistan, all the subject countries switch their scheme of sectoral classifications from ISIC-2 to ISIC-3. The shift in classification occurs at different points in time for different countries. Furthermore, the ISIC classification is somewhat different from the NAICS classification between the years 2002 till 2013 which is followed by the reference country U.S. However, this can be adjusted with simple data transformations. TABLE 3.4 summarizes all such transformations I make to present sectoral data consistently across varying schemes of ISIC-2, ISIC-3 and NAICS classifications.

(ii) Sectoral Division of Employment:

Employment data is sourced from two distinct databases of International Labour Organization (ILO) published by ILO Department of Statistics. Annual data from the year 1970 to 2008 is obtained from LABORSTA and from year 2009 to 2013 the data is taken from ILOSTAT. Similar to output data, the sectoral employment data provided by ILO is also not consistent over years or across countries. But, broadly, the sectors match with the seven sectors suggested by UNCTAD[3].

[3] ILO also follows International Standard Industrial Classification (ISIC) for categorizing sectors.

TABLE 3.3: Scheme of Sectoral Division (Value Added) Varying over Years and Across Countries

Country	ISIC-Revision.2	ISIC-Revision.3	NAICS
Hong Kong (1978-2008)	1978-2008	-	-
Indonesia (1976-2008)	1976-1999	2000-2008	-
Japan (1970-2013)	1970-2002	2003-2013	-
Korea (1970-2013)	1970-1991	1992-2013	-
Malaysia (1980-2013)	1980-2000	2001-2013	-
Pakistan (1973-2013)	1973-2008	-	-
Philippines (1971-2013)	1971-2000	2001-2013	-
Singapore (1970-2006)	1970-1984	1985-2006	-
Sri Lanka (1981-2010)	1981-2001	2002-2010	-
Thailand (1971-2013)	1971-2001	2002-2013	-
United States (1970-2013)	1970-2001	-	2002-2013

TABLE 3.4: Transformation of ISIC and NAICS Sectoral Output Schemes to Sectoral Divisions (7 and 4 Sector Groups)

Sector of the Economy	*ISIC-2 Disaggregation into 7 Sectors*	*ISIC-2 Disaggregation into 4 Sectors*	*ISIC-3 Disaggregation into 7 Sectors*	*ISIC-3 Disaggregation into 4 Sectors*	*NAISC-2007 Disaggregation into 7 Sectors*	*NAISC-2007 Disaggregation into 4 Sectors*
Agriculture, Hunting, Forestry and Fishing	1 Division 11-13.	1 Agriculture	A+B Divisions 01-05	A+B Agriculture	Division 11	11 Agriculture
Manufacturing	3 Divisions 31-39.	3 Manufacturing	D Divisions 15-37	D Manufacturing	Division 31-33	31-33 Manufacturing
Mining, Quarrying and Utilities	2+4 Division 21-29 & 41-42	2+4+5 Industry	C+E Divisions 10-14& 40-41	C+E+F Industry	Division 21 & 22	21-23 Industry
Construction	5 Division 50		F Division 45		Division 23	
Wholesale, Retail Trade, Restaurants & Hotels	6 Divisions 61-63	6+7+8+9 Services	G+H Divisions 50-55	G-P Services	Division 42	42-92 Services
Transport, Storage and Communications	7 Divisions 71-72		I Divisions 60-64.		Division 48-49 & 51	
Other Activities	8+9 Divisions 81-96		J-P[4] Divisions 65-95.		Division 52-92	

[4] See footnote 5 for details.

As mentioned in the previous section, ILO offers sectoral employment data with North American Industry Classification System (NAICS) for the reference country U.S. from 2002 to 2013. The sectoral division suggested by NAICS does not exactly match with that of ISIC. This serves as another potential source of discrepancy in the data[5]. The sectors are classified at finer (more disaggregated) levels relative to ISIC revision 2 and 3. NAICS (1997) United States in its original version has 20 sectors, 96 subsectors, 311 industry groups and 1175 U.S. industries. In comparison, ISIC has 17 sections, 60 divisions, 150 groups and 292 classes. The 2-digit sectors remained unchanged through the four revisions of classification (NAICS 1997 TO NAICS 012); however, all the subsectors, industrial groups and final industries are continuously being updated from one revision to other. We eliminate the apparent incomparability by aggregating finely classified sectors under the relevant main sector so that sectoral employment could be reasonably matched. TABLE 3.6 to 3.8 explain how the employment data of Asia and U.S. under ISIC and NAICS classifications is transformed to construct sectoral employment data comparable with my 7-sector and 4-sector output divisions.

[5] This mis-match in classification is an issue with all studies involving U.S. data from 2002 – 2013. But the literature has never explained how this has been resolved.

TABLE 3.5: Sectoral Disaggregation of Real Economy followed by UNCTAD

Top Level Classification Codes	Sector of the Economy	Aggregation into Seven Sectors (ISIC-Revision.3)	Aggregation into Four Sectors
A-B	Agriculture, Hunting, Forestry and Fishing	(A-B) The sector corresponds to ISIC Rev.3 div 01-05.	*Agriculture (A-B)*: Includes agriculture, hunting, forestry and fishing. It corresponds to ISIC Rev.3 divisions 01-05.
C &E	Mining, Quarrying and Utilities	(C & E) The sector corresponds to ISIC Rev.3 div 10-14, 40 and 41.	*Industry (C, E & F)*: Includes mining and quarrying, electricity, gas and water supply, and construction. It corresponds to ISIC Rev.3 divisions 10-14, 40-45.
F	Construction	(F) The sector corresponds to ISIC Rev.3 div 45.	
D	Manufacturing	(D) The sector corresponds to ISIC Rev.3 div 15-37.	*Manufacturing (D)*: It corresponds to ISIC Rev.3 divisions 15-37.
G-H	Wholesale, Retail Trade, Restaurants and Hotels	(G-H) The sector corresponds to ISIC Rev.3 div50-55.	*Services (G-P)*: Include all other economic activities (it corresponds to ISIC Rev.3 divisions 50-55 and 60-95.
I	Transport, Storage and Communications	(I) The sector corresponds to ISIC Rev.3 div 60-64.	
J-P	Other Activities	(J-P) The sector corresponds to ISIC Rev.3 div 65-95.	

TABLE 3.6: Sectoral Disaggregation of Total Employment followed by ILO (ISIC-Revision 2)

ISIC-Revision 2	Aggregating Employment into Seven Sectors	Aggregating Employment into Four Sectors
1. Agriculture, Hunting, Forestry and Fishing	1. Agriculture, Hunting, Forestry and Fishing	*Agriculture (1)*: Includes agriculture, hunting, forestry and fishing.
2. Mining and Quarrying	2+4. Mining and Quarrying, Electricity, Gas and Water	
3. Manufacturing	3. Manufacturing	*Industry (2+4+5)*: Includes mining and quarrying, electricity, gas and water supply, and construction.
4. Electricity, Gas and Water		
5. Construction	5. Construction	*Manufacturing (3)*: Includes manufacturing only.
6. Wholesale and Retail Trade and Restaurants and Hotels	6. Wholesale, Retail Trade, Restaurants and Hotels	
7. Transport, Storage and Communication	7. Transport, Storage and Communication	*Services (6+7+8+9)*: Includes wholesale, retail trade, restaurants and hotels, transport, storage and communication, financing, insurance, real estate and business services, community, social and personal services
8. Financing, Insurance, Real Estate and Business Services	8+9. Other Activities	
9. Community, Social and Personal Services		

TABLE 3.7: Sectoral Disaggregation of Total Employment followed by ILO (ISIC-Revision.3)

ISIC-Revision.3	Aggregating Employment into Seven Sectors	Aggregating Employment into Four Sectors
A-B. Agriculture, hunting, forestry and fishing	(A-B)-Agriculture, Hunting, Forestry and Fishing	*<u>Agriculture (A-B)</u>*: Includes agriculture, hunting, forestry and fishing.
C. Mining and Quarrying	(C+E)-Mining and Quarrying, Electricity, Gas and Water	*<u>Industry (C+E+F)</u>*: Includes mining and quarrying, electricity, gas and water supply, and construction.
D. Manufacturing	(F)-Construction	
E. Electricity, Gas and Water Supply		
F. Construction	(D)-Manufacturing	*<u>Manufacturing (D)</u>*: Includes manufacturing only.
G. Wholesale and Retail Trade; Repair of Motor Vehicles, Personal and Household Goods	(G+H)-Wholesale, Retail Trade, Restaurants and Hotels	*<u>Services (G-P)</u>*: Includes wholesale, retail trade, restaurants and hotels, transport, storage and communication, financing, insurance, real estate and business services, community, public administration, education, health and social services.
H. Hotels and Restaurants		
I. Transport, Storage and Communications	(I)-Transport, Storage and Communication	
J-L & O-P[6]	(J-P)- Other Activities	
M. Education		
N. Health and Social Work		

[6] Financial intermediation, real estate, renting and business activities, public administration and defence; compulsory social security, other community, social and personal service activities, private households with employed persons.

TABLE 3.8: Sectoral Disaggregation of Total Employment followed by NAICS (2007)

NAICS (2007)		Aggregating Employment into Seven Sectors	Aggregating Employment into Four Sectors
11. Agriculture, Forestry, Fishing and Hunting	53. Real Estate and Rental and Leasing	(11)-Agriculture, Hunting, Forestry and Fishing	*Agriculture (11)*: Includes agriculture, hunting, forestry and fishing.
21. Mining, Quarrying, and Oil and Gas Extraction	54. Professional, Scientific, and Technical Services	(21-22)-Mining and Quarrying, Electricity, Gas and Water	*Industry (21-22+23)*: Includes mining and quarrying, electricity, gas and water supply, and construction.
22. Utilities	55. Management of Companies and Enterprises		
23. Construction	56. Administrative and Support and Waste Management and Remediation Services	(23)-Construction	
31-33. Manufacturing	61. Educational Services	(31-33)-Manufacturing	*Manufacturing (31-33)*: Includes manufacturing only.
42. Wholesale Trade	62. Health Care and Social Assistance	(42+44-45)-Wholesale, Retail Trade, Restaurants and Hotels	*Services (42-92)*: Includes wholesale, retail trade, restaurants and hotels, transport, storage and communication, financing, insurance, real estate and business services, community, public administration, education, health, professional, scientific, and technical services, administrative and support and waste management and remediation services, accommodation and food services and social services.
44-45. Retail Trade	71. Arts, Entertainment, and Recreation		
48-49. Transportation and Warehousing	72. Accommodation and Food Services	(48-49+51)-Transport, Storage and Communication	
51. Information	81. Other Services (except Public Administration)	(52-92)-Other Activities	
52. Finance and Insurance	92. Public Administration		

3.2.2 Classification of Traded and Non-Traded Sectors

Tradability (nontradability) of goods and services is a much debated feature of the real economy. Though researchers suggest a couple of methods for distinguishing between traded and nontraded sectors, these methods are not well-acknowledged as they are ad hoc in nature. Majority of the studies on BS hypothesis distinguish between traded and no-traded sectors arbitrarily (see Canzoneri et al., 1999; Chinn, 2000; Egert et al., 2003). Some studies rely on earlier studies (analysing similar set of countries) for differentiating between tradable and nontradable sectors. For example, Kakkar (2002, 2003) follows the sectoral division of real economy into tradables and nontradables as suggested by De Gregorio, Giovannini and Wolf (1994) and Stockman and Tesar (1995). Similarly, Thomas and King (2008) empirically test BS hypothesis for Asia-Pacific countries by adopting Chinn's (2000) sectoral classification that tests the hypothesis for the same region.

In this dissertation, I follow the sectoral division suggested by Dumrongrittikul (2012). He estimates BS model empirically for a set of 33 developing and developed economies. Out of these 33 countries, 14 are from East Asia, South East Asia and South Asia, the regions I investigate in my analysis. All the countries that I study are present in Dumrongrittikul's (2012) sample with the sole exception of Hong Kong. His data set ranges from the year 1970 to 2008, but I extend the data through 2013 for majority of my sample countries (see TABLE 3.3 for sample periods for each country).

I follow Dumrongrittikul's (2012) study as it classifies sectors using a robust approach. First, it uses the traditional method of sectoral classification as suggested by De Gregorio et.al (1994)[7]. The study also employs a second method of cointegration for differentiating between

[7] De Gregorio, Giovannini and Wolf (1994) consider an industry as tradable if the ratio of its exports to total production volume is greater than 10 percent.

tradable and nontradable sectors of an economy. The second approach is originally proposed by Gonzalez-Soriano (1990)[8]. The author combines these two approaches to establish a criterion for classifying the traded and nontraded sectors. The industry that yields a tradability ratio[9] of less than 0.1 is treated as nontradable. The industry that yields a tradability ratio of higher than 0.2 is treated as tradable. If the tradability ratio lies between 0.1 and 0.2, price co-movement test is conducted to decide the tradability or nontradability of the industry. Using two tests simultaneously minimize the chances of error in the sectoral classification.

I also see Dumrongrittikul's sectoral classification strategy as an improvement over pre-existing approaches to sectoral division for another critical reason. A vast number of BS studies follow a uniform classification of industries between tradables and nontradables across their set of sample countries. This is an unrealistic assumption of industry homogeneity (with respect to its tradability in international markets) that has been imposed in the earlier studies. Such a practice may prevent the researcher from capturing the actual degree of tradability in each country which may result in misleading representation of sectors whilst constructing price and productivity series. However, Dumrongrittikul's suggested scheme of sectoral breakup extends a fairly convincing solution to this discrepancy and allows country specific heterogeneity over each industry. Thus, by allowing sectoral weights and industry heterogeneity in the construction of my price and productivity series, I account for changing composition of traded and nontraded sectors in both cross-sectional and time dimensions.

[8] The method devised by Gonzalez-Soriano (1990) takes long-run convergence of inter-country tradable prices (i.e., the valid existence of PPP) as an evidence in support of tradability of an industry. Using Engle-Granger (1987) method of cointegration, the author verifies the tendency of long-run co-movement between the home and foreign prices of each of sector to decide about its tradability (nontradability).

[9] Tradability of a certain industry is measured by the ratio of its exports plus imports to total industrial output.

3.2.3 Hong Kong - A Special Case:

Hong Kong is the only country in my analysis whose sectoral division does not follow Dumrongrittikul (2012) as it was not a part of his analysis. In order to classify Hong Kong's economy into traded and nontraded sectors, I follow Li (2005). The author evaluates Hong Kong's competitiveness through its sectoral total factor productivity by dividing the real economy into three distinct sectors: tradable goods sector, tradable services sector and nontradable sector. However, Li (2005) classification does not map one-to-one to the seven sector classification of my study. Li employs a more disaggregated sectoral division than mine. His classification divides Hong Kong's real economy into twenty distinct sectors. As a result, many of the industries which are practically inseparable in my analysis are distinctively separated in Li's data. For example, mining, quarrying and utilities represent a single sector in my data, but constitute two distinct sectors in Li (2005) data. Mining and quarrying are categorized under traded sector and utilities is treated as nontradable. Same is true for wholesale, retail trade, restaurants and hotel industry which is taken as a single sector in my analysis but is treated as two differentiated sectors by Li (2005). Hotel industry is categorized as tradable service and wholesale, retail trade and restaurants are classified as nontradable sector[10]. Thus, in my analysis, a one-to-one mapping of sectors as tradables and nontradables based on Li (2005) is not possible. However, for robustness, I adopt four possible sectoral classifications for Hong Kong based on Li (2005) in my analysis. Each type is individually outlined below.

Classification I (HKG_1): The first classification treats construction as the only nontradable sector of the economy. Rest of the sectors are treated as tradables.

[10] See Li (2005) for details on the different sectoral classifications for Hong Kong.

Classification II (HKG_2): The second classification adds both mining, quarrying & utilities, and wholesale, retail trade, restaurants & hotel industry to the nontradable sectors together with construction. Agriculture, manufacturing, transport, storage & communications and other activities are categorized as traded sectors.

Classification III (HKG_3): The third classification has construction and wholesale, retail trade, restaurants & hotel industry as nontradable sectors and agriculture, mining, quarrying & utilities, manufacturing, transport, storage & communications and other activities as tradable sector.

Classification IV (HKG_4): Finally, I have mining, quarrying & utilities and construction as nontraded sectors and agriculture, manufacturing, wholesale, retail trade, restaurants & hotel industry, transport, storage & communications and other activities as traded sectors.

The classification of countries' different sectors into traded and nontraded are presented in Table 3.9 and 3.10 below.

TABLE 3.9: Dividing Real Economy into Traded and Non-Traded Sectors as Proposed by Dumrongrittikul (2012)

Country	Agriculture, Hunting, Forestry & Fishing	Manufacturing	Mining, Quarrying & Utilities	Construction	Wholesale, Retail Trade, Restaurants & Hotels	Transport, Storage & Communications	Other Activities
Indonesia (1976-2013)	NT	T	T	NT	T	T	NT
Japan (1970-2013)	NT	T	NT	NT	NT	T	NT
Korea (1970-2013)	NT	T	T	NT	T	T	NT
Malaysia (1980-2013)	T	T	T	T	NT	T	T
Pakistan (1973-2008)	T	T	T	NT	NT	T	NT
Philippines (1971-2013)	NT	T	T	NT	NT	T	NT
Singapore (1970-2006)	T	T	T	NT	T	T	T
Sri Lanka (1981-2010)	T	T	NT	NT	NT	T	NT
Thailand (1971-2013)	NT	T	T	NT	NT	NT	NT
United States (1970-2013)	T	T	NT	NT	NT	NT	NT

Disaggregation into 4 Categories: (Agriculture, Hunting, Forestry and Fishing) + Manufacturing = **T**
(Uniformly applies to all countries): Industry[(Mining, Quarrying & Utilities) + Construction] + Services [(W. Sale, R. Trade & Hospitality)+Transport & Communication +Other Activities] = **NT**

TABLE 3.10: Sectoral Classification for Hong Kong

Sectors of the Economy	Classification I HKG_1	Classification II HKG_2	Classification III HKG_3	Classification IV HKG_4
Agriculture, Hunting, Forestry and Fishing	T	T	T	T
Manufacturing	T	T	T	T
Mining, Quarrying and Utilities	T	NT	T	NT
Construction	NT	NT	NT	NT
Wholesale, Retail Trade, Restaurants & Hotels	T	NT	NT	T
Transport, Storage and Communications	T	T	T	T
Other Activities	T	T	T	T

3.3 A First Look at the Data: Sectoral Prices and Productivity

In this section, I look at the evolution of the sectoral prices and productivity series over time for all countries. Observing the trend and fluctuations of prices and productivity series are helpful in obtaining a preliminary (though informal) insight into the conceivable long-run relationship between sectoral prices and productivity.

FIGURE 3.1 shows the time plots for prices and productivity in two individual panels. For each country, the first panel (left hand side) contains internal sectoral prices of tradables and nontradables. The right hand side panel displays the internal sectoral productivities. The price deflators and the average labour productivity involved in the composition of two types of series follow the 7-sector division.

I first look at the time series plots for Indonesia. The sectoral prices are sharply trending upwards. Nontradable sector price movements have always been greater than or equal to tradable prices. An abrupt jump in sectoral prices can be seen from the years 1997-2001. Starting from the period of East-Asian financial crisis a phase of historical decline in country's growth and intense inflation followed. The sectoral productivities are trending at a modest rate over the sample period. During the first eight years of the sample period, the traded sector productivity initially falls dramatically and then trends upwards. Such a fluctuation can be attributed towards oil prices hike after 1973 making the government adopt inward-looking and restrictive trade, investment and business policies.

I next look into Japan's prices and productivity data. From the year 1970 till 1990, there is an upward trend of sectoral prices which is well explained through the country's monetary expansion pursed in 1973. Also the surging stock and land prices and the outbreak of war in the Middle East led to an oil crisis. However, after the burst of nation-wide asset price bubble, the country faced a situation of severe deflation (real estate and construction were affected most) which can be seen clearly through the falling sectoral prices after the year 1990. The traded sector productivity is out pacing its counterpart for almost the entire data period. This sharp rise in productivity of tradable sector is contributed by rampant growth in country's manufacturing sector where the level of productivity has always been higher than in other sectors.

The case of Korea is very interesting. The time plot of productivity in tradables is visibly departing away from nontradables one with sizeable growth. This implies rising trend in tradables productivity, whereas nontradables productivity tend to remain rather stagnant over the sample period. On part of sectoral prices, both traded and nontraded sector prices have sharply inflated over the time. From 1970 to 2013, this rise has been of more than 100 percent.

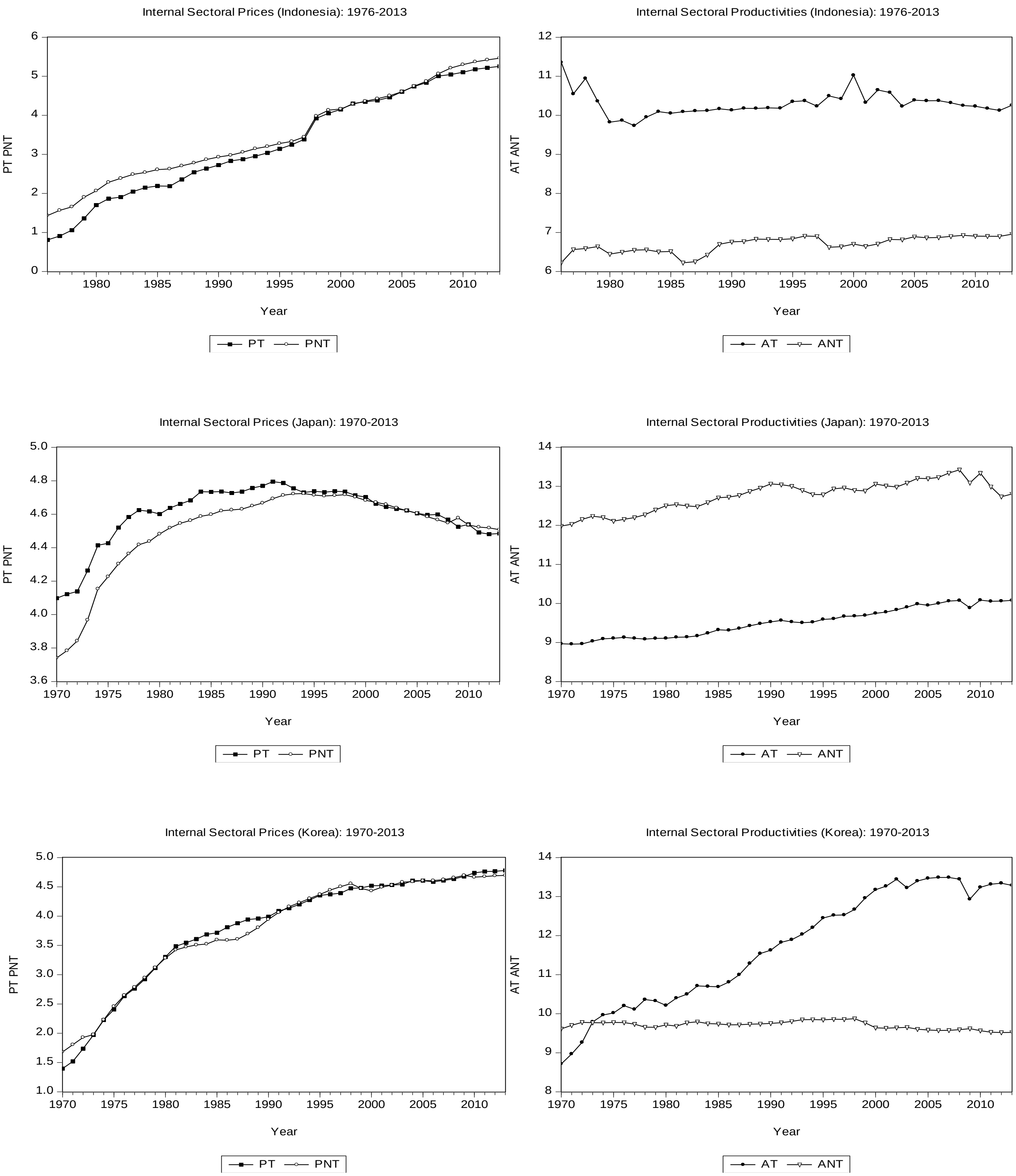
The time series of sectoral prices and productivity for Malaysia and Philippines are quite similar to that of Indonesia. The sectoral prices are sharply rising over the sample period, with Philippines reporting an immense growth rate of prices. The post Asian financial crisis inflation in Philippines is self-explanatory. The country was one of most adversely affected states in East Asia. However, the pre-crisis inflation is dominantly explained by adjustment to the balance of payments and short-term external debt crisis in 70s, the oil price hike of 1973, public sector expansion and sharp devaluation of peso, country-wide energy crisis and sizeable monetary expansion in the first half of 90s. The real side of the economy is not performing very well. Both traded and nontraded sector productivity do not seem to be growing at a sufficiently faster rate. Rather the two series tend to fluctuate around their mean value displaying no significant trend movements.

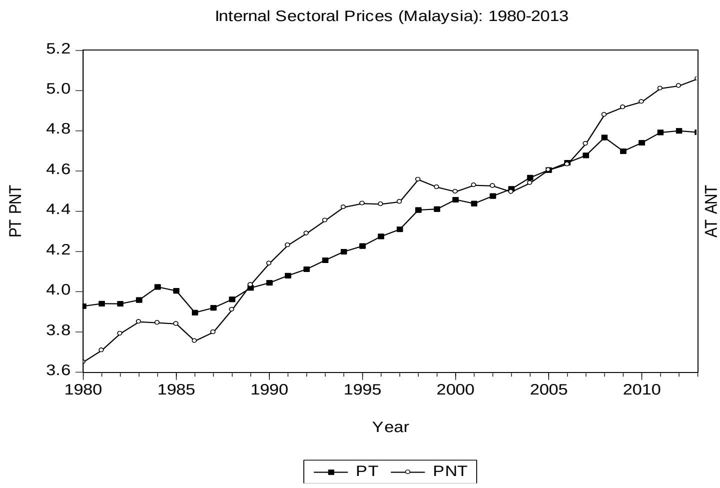
Pakistan and Sri Lanka are the two countries from the South Asian region. From the visual inspection, the sectoral prices demonstrate a sharp rise over the sample period and the sectoral productivity growth performance has also been significant during this time. One noticeable feature on part of sectoral productivity is that the tradable sector productivity is growing at an exceeding rate (relative to nontradable sector productivity). This is particularly true for Pakistan. The behaviour is well-evident from South Asian merchandise trade which, as percentage of GDP, has grown from 13.4 percent in 1960 to 31.7 percent in 2015.

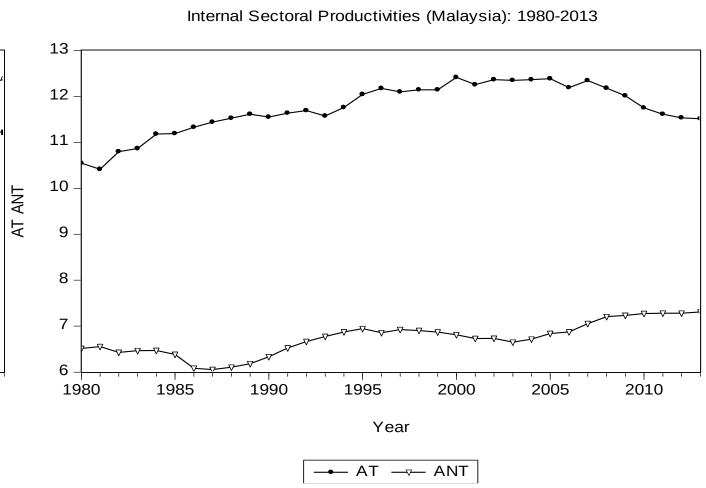
Singapore's nontradable sector prices and productivity does not have a clear trend during my sample period. The series sharply trends upward initially (year 1970 to 1982) but afterwards displays a fluctuating behaviour with large periodical swings. Singapore's tradable sector productivity and prices tend to grow at a modest pace over time.

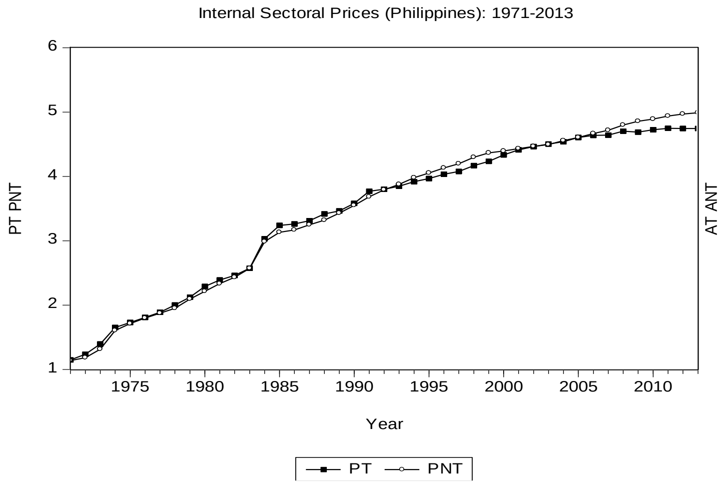
Finally, looking into the sectoral prices and productivity time plots of Thailand from 1970 to 2013, I find that sectoral prices tend to rise sharply before 1997 but become relatively constant afterwards. Also, the growth in traded sector productivity, largely driven by the immense growth of the country's industrial sector, has been sizeable throughout my sample period, and clearly surpasses the nontradable sector productivity growth performance.

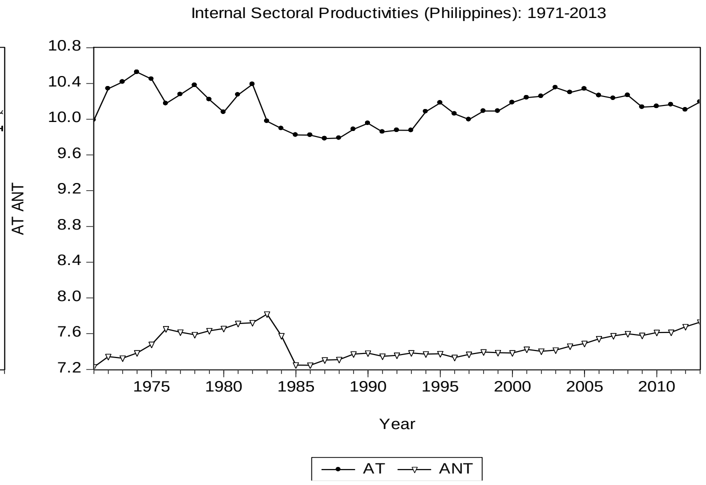
FIGURE 3.1: Internal Prices and Productivity Gaps of Subject Asian Countries (Seven Sectors)

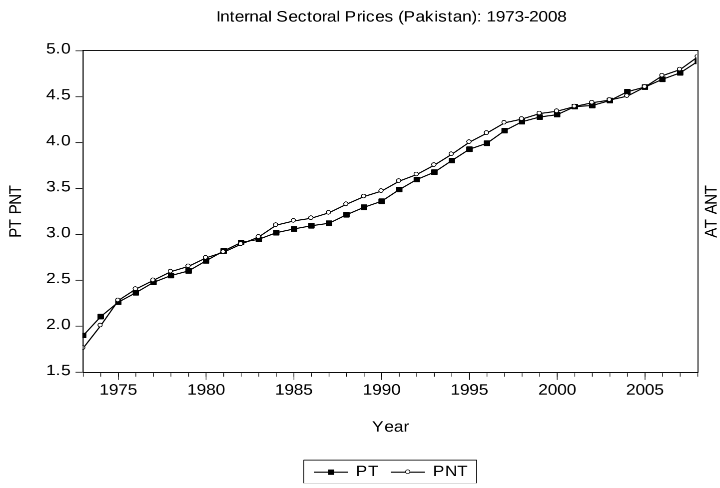


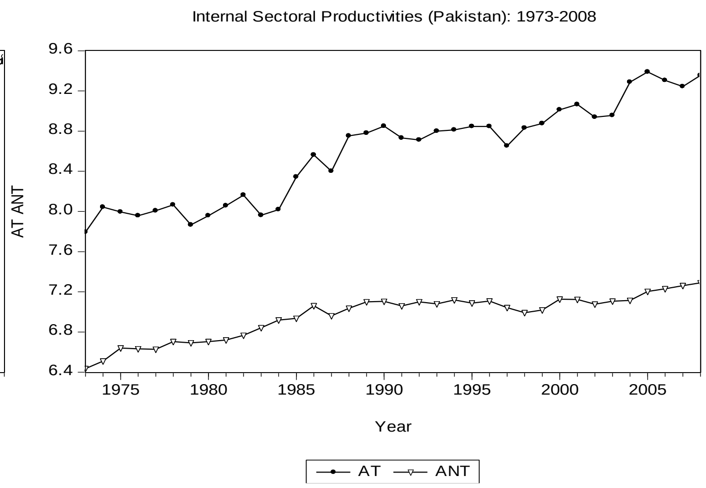

Internal Sectoral Prices (Malaysia): 1980-2013
PT PNT
Year
PT PNT


Internal Sectoral Productivities (Malaysia): 1980-2013
AT ANT
Year
AT ANT


Internal Sectoral Prices (Philippines): 1971-2013
PT PNT
Year
PT PNT


Internal Sectoral Productivities (Philippines): 1971-2013
AT ANT
Year
AT ANT


Internal Sectoral Prices (Pakistan): 1973-2008
PT PNT
Year
PT PNT


Internal Sectoral Productivities (Pakistan): 1973-2008
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Year
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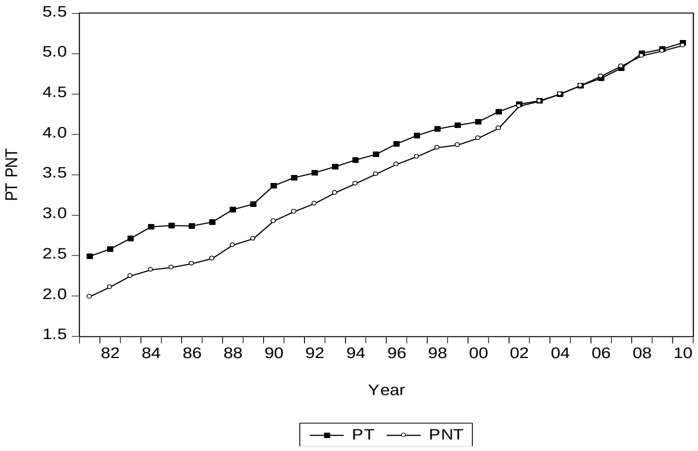
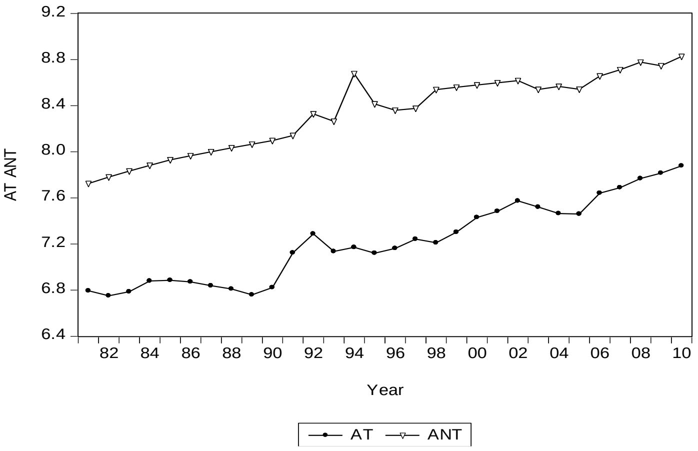
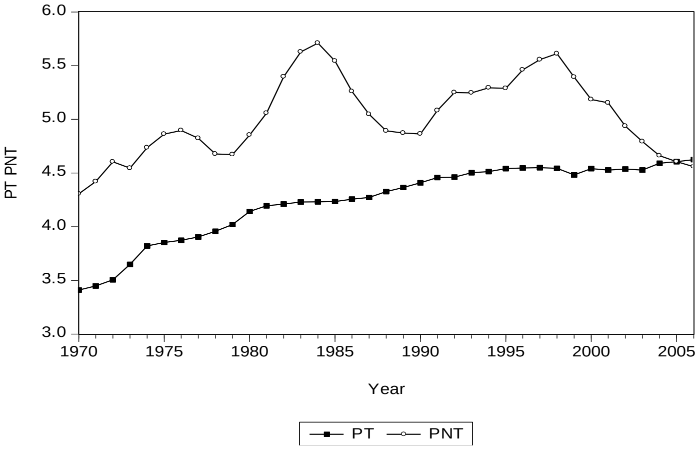
Internal Sectoral Prices (Sri Lanka): 1981-2010



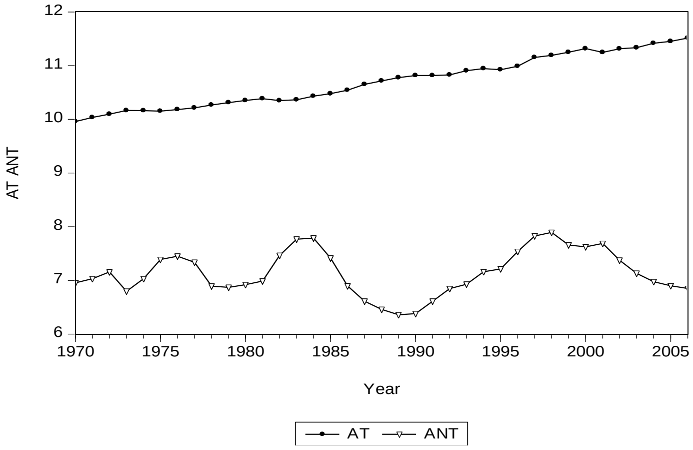
Internal Sectoral Productivities (Sri Lanka): 1981-2010



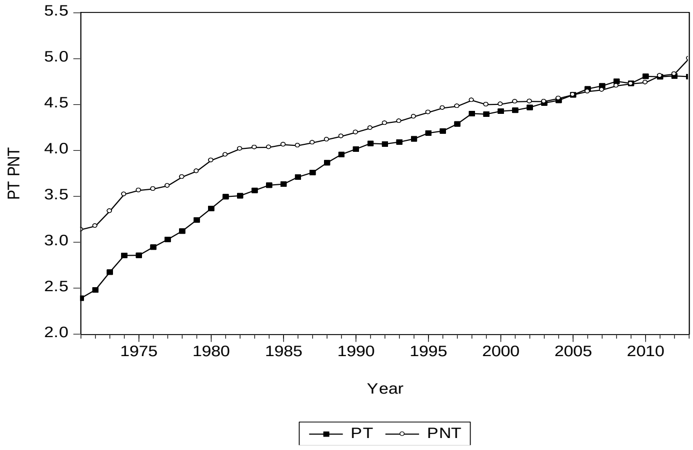
Internal Sectoral Prices (Singapore): 1970-2006



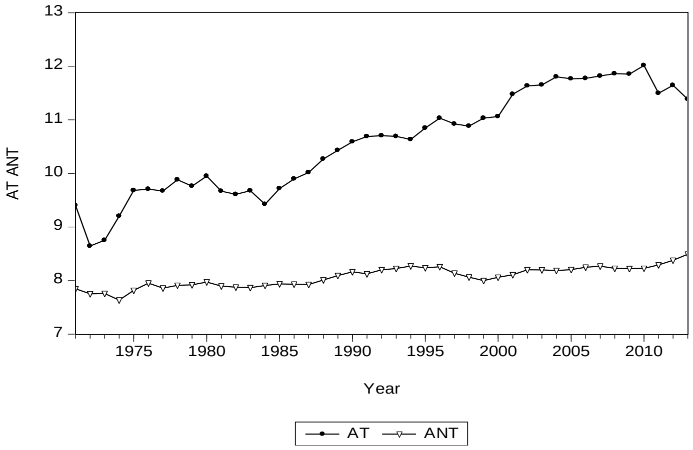
Internal Sectoral Productivities (Singapore): 1970-2006



Internal Sectoral Prices (Thailand): 1971-2013



Internal Sectoral Productivities (Thailand): 1971-2013



CHAPTER 4: REVIEW OF THE LITERATURE

4.1 Background

The aim of this chapter is to provide a detailed insight into the existing empirical work done on the long-run real exchange rate behaviour in the context of the Balassa-Samuelson hypothesis. Since the literature on this topic is quite vast, I will primarily focus on the empirical studies that investigate the hypothesis in the context of the Asian economies.

The empirical studies on Asia that conduct a multi-country examination of the Balassa-Samuelson hypothesis are inconsistent in dealing with different features of the data like sectoral disaggregation; definitions of the real exchange rate and prices; consistency of output and employment series; and empirical methodology. Hence the results from these countries are very mixed, and in many cases, not robust. In contrast, a sizeable number of studies on the Balassa-Samuelson hypothesis investigating countries in Europe and OECD are insightful as they explain and verify many of the aspects of the hypothesis which are critically important in yielding reliable model estimates. These studies have carefully considered the issues such as coverage of industries and scheme of industrial classification between tradables and non-tradables (Rother, 2000; Mihajljek et al., 2003 & 2004; Coricelli and Jazbec, 2004; Egert, 2003, 2004 & 2005; Gibson, 2008); choice of real exchange rate, price and productivity measures (Canzoneri et al. 1999; Egert et al. 2003; Lee and Tang, 2007); set of empirical estimation methods applied (Chong et al., 2012; Boreo et al., 2012); theoretical specifications of the hypothesis and testing the model assumptions (Strauss and Ferris, 1996; Estrada and Lopez-Salido, 2004; MacDonald and Ricci, 2001, 2005;

Schmillen, 2013) and revealed how sensitive the empirical estimates of the hypothesis are to correctly addressing these issues. In light of these considerations, I lay down the following four criteria to gauge the relative performance of the studies analysing the Balassa-Samuelson hypothesis for Asia. The primary objective of establishing these evaluation parameters is to obtain a comprehensive guideline to critically evaluate the existing studies and empirically test the Balassa-Samuelson model and provide consistent and robust model estimates. The criteria are:

(a) Measurement of real exchange rate

(b) Scheme of sectoral classification

(c) Empirical estimation techniques

(d) Theoretical foundations of the model: This is further categorized into:

  (i) Domestic version of BS hypothesis

  (ii) Assumptions of the Balassa-Samuelson model

In the forthcoming sections of the chapter, I discuss each of these parameters and provide a critical review of the different aspects in the context of empirical studies on Asia.

4.2 Evaluating the Existing Literature on Established Parameters

4.2.1 Measurement of Real Exchange Rate

Empirical verification of the Balassa-Samuelson hypothesis necessitates the precise measurement of real exchange rate for obtaining consistent model estimates. However, the task is much complicated due to the lack of consensus by researchers on the most reflective real exchange rate proxy that is capable of reflecting the internal transmission channels from domestic prices to real exchange rate.

In the empirical studies on Asia, I find that CPI is the most popular price deflator used to construct the real exchange rate. Drine and Rault (2002) use CPI based effective real exchange rates of six Asian economies to test the Balassa-Samuelson hypothesis. Wang and Dunne (2003) verify the Balassa-Samuelson effect for a group of seven East Asian countries using CPI based bilateral real exchange rates on quarterly data. Choudhri and Khan (2005) explain long-run behaviour of CPI based bilateral real exchange rates for a panel data set of 16 developing countries from East and South Asia, Western Hemisphere, and Africa. Tsen (2011), Dumrongrittikul (2012), Kakkar and Yan (2012), Ricci et al. (2013) and Wang et al. (2016) also verify the Balassa-Samuelson hypothesis by using CPI deflated bilateral real exchange rates for emerging Asia, in addition to Ex-Asian developing and developed countries.

However, the index in not without its limitations. Firstly, CPI is highly exposed to price controls, subsidies and indirect taxes which may distort the role of market forces in determining prices (see Hinkel & Montiel (1999)). Secondly, CPI may have a fairly large number of imported items in its basket. Thirdly, degree of weights involved in the construction of CPI across countries can vary quite a bit. This causes a hindrance in cross-country comparisons of cost effectiveness and competitiveness. In such a situation, a rise in price of a certain commodity may mislead towards improvement or deterioration in relative competitiveness between countries. The problem is more visible in developing countries which seriously lack a representative sample of goods and services in their CPI. Thus, despite its wide popularity as a measure of non-tradable prices, the measure holds some important caveats.

Some important studies analyse the Balassa-Samuelson hypothesis for Asia using national output deflator (GDP or GNP) based real exchange rate (Bahmani-Oskooee and Rhee, 1996; Ito et al., 1999; Chinn, 2000; Bahmani-Oskooee and Nasir, 2004; Thomas and King, 2008). However, GDP deflator as the national price index is also not the most perfect measure. This measure has been criticized in literature since GDP deflators do not necessarily correspond to the officially published inflation indices, which are normally represented by CPI, PPI or WPI rather than output deflators. Moreover, for modelling relative price movements between two countries, base-weighted price index is preferred over a current-weighted price index (Goldstein et al, 1980).

In my thesis, I employ three different price indices – CPI, GDP deflator, and GDP deflator using non-tradable sector prices only – to construct three distinct variants of real exchange rate. Employing three different measures of real exchange rate is useful as (a) it will reveal the relative performance of three alternative real exchange rate measures, and (b) it checks the robustness and consistency of my empirical estimates for the three alternative versions of the Balassa-Samuelson model (please refer to chapters five and six).

4.2.2 Scheme of Sectoral Classification

The correct distinction between traded and non-traded sectors of an economy is crucial for the empirical verification of the Balassa-Samuelson hypothesis. Nevertheless, there is a serious lack of consensus on recognizing sectors as tradables or non-tradables owing to the disagreement at the conceptual level for measuring tradability (non-tradability) of an industry. The problem is further aggravated when the level of aggregation of the existing data may be too high to permit a clear classification of industries into one sector or the other. This concern goes largely unaddressed in the existing empirical literature on Asia. A vast majority of the studies analysing countries in Asia classify sectors in a rather subjective manner. Such a practice casts shadow over the reliability of their empirical estimates.

Ito et al. (1999) examine the Balassa-Samuelson hypothesis for Asia Pacific Economic Cooperation (APEC) member states and the Western Hemisphere and Oceania regions. The real economy of each country is classified into traded and non-traded sectors by assuming manufacturing sector as tradable and services as non-tradable. The distinction is made completely arbitrarily under the belief that East Asian productivity growth is substantially driven by their high value-added manufactured goods exports (e.g. machine exports). Another important drawback of their study is that the coverage of sectors is highly aggregated which may result in biased estimates of sectoral productivity growth on relative price movements. A number of prominent studies on the Balassa-Samuelson hypothesis for Europe and other Ex-Asian regions recognize the importance of more formal methods of sectoral division and better industry coverage for obtaining reliable model estimates (see Rother, 2000; Mihajljek et al.,2003, 2004; Coricelli and Jazbec, 2004; Egert, 2004).

Choudhri and Khan (2005) covers a set of sixteen developing countries in Asian, African and Western states with considerably varying economic structures. However, each of them is subject to a standard sectoral division with agriculture and manufacturing as tradable sectors and rest of the economy as non-tradables. For empirical verification of the Balassa-Samuelson hypothesis, the sectoral output and price data are sourced from World Development Bank Indicators (WDI) database. The database contains sectoral value-added and price deflator time-series but not with fine sectoral classification. Each country is disaggregated into four distinct sectors - agriculture, manufacturing, industry and services - and these sectors are too broad to capture the internal transmission mechanism of the model. Using such a highly aggregated data may result in forced (and even wrong) assignment of certain industries as non-tradables, whereas the industries should actually be classified as tradables (or at least partly tradables) due being exposed to foreign competition. The classic examples are the air and ship transportation services, distribution sector and the utilities industries which are normally treated as non-tradables (see MacDonald and Ricci, 2005; Thomas and King, 2008).

In addition to above discussed studies, a vast majority of Asian studies distinguish between traded and non-traded sectors arbitrarily (Bahmani-Oskooee and Rhee, 1996; Chinn, 2000; Wang and Dunne, 2003; Bahmani-Oskooee and Nasir, 2004; Drine and Rault, 2004; Olson, 2009;

Chowdhury, 2012). Some studies rely on earlier studies (analysing similar set of countries or even dissimilar) for deciding between the tradability (non-tradability) of sectors. For example, Thomas and King (2008) empirically test the Balassa-Samuelson hypothesis for Asia-Pacific countries by following Chinn's (2000) sectoral classification who tests the model for the same region. Similarly, Kakkar (2012) analyse six East Asian countries following the sectoral division suggested by De Gregorio, Giovannini and Wolf (1994) and OECD countries following Stockman and Tesar (1995). Similarly, Ricci et.al (2013) follow De Gregorio, Giovannini, and Wolf (1994) sectoral classification for empirically investigating the augmented version of the Balassa-Samuelson hypothesis for a set of 48 industrial and emerging economies (including East and South Asian regions). In practice, adopting the sectoral division of earlier studies is acceptable provided (a) the base study covers the similar set of countries (or countries with reasonably similar economic structures), and (b) the base study categorizes sectors using some formal and well-recognized method of sectoral division rather than distinguishing sectors in a purely subjective way.

The only study on Asia that stands out in the literature in the context of precise sectoral classification is done by Dumrongrittikul (2012). The study is distinct since the author uses a combination of two approaches for sectoral division. The study tests the Balassa-Samuelson hypothesis empirically for a set of 33 developed and developing economies (including 14 countries from East and South Asia). The sectors are divided at a sufficiently disaggregated level into seven distinct industries. The study is distinctive in the sense that the author employs the Engle-Granger (1987) single equation error correction model in addition to the traditional method of calculating tradability ratio. The method allows for country-specific heterogeneity over each industry and changes in classification along the period. The basic belief behind this method is that tradable commodities across countries are likely to satisfy the law of one price (LOOP) and purchasing power parity (PPP). The author suggests estimating the 2-step Engle-Granger error correction model by regressing the domestic price level of each sector on the corresponding sector's price in the U.S, the reference country in his analysis. Rejection of the null hypothesis at a desirable significance level displays the potency of the domestic sectoral price series to co-move with the international market prices establishing PPP. This allows for the distinction between tradable and non-tradable sectors.

In my dissertation, I test the Balassa-Samuelson hypothesis empirically by adopting two different approaches to sectoral division: (a) dissecting the real economy into four distinct sectors and classifying traded and nontraded sectors somewhat arbitrarily following some of the previous literature (see chapters three and five for detailed notes on variable definitions, data source, sectoral division and empirical results), and (b) following a more precise and reliable seven sector classification as suggested by Dumrongrittikul (2012) (see chapters three and six for detailed notes on variables definitions, data source, sectoral division and empirical results). Estimating the Balassa-Samuelson hypothesis for two different data sets with alternative schemes of sectoral division will serve as a robustness check for my model estimates.

4.2.3 Empirical Estimation Techniques

Estimates of the Balassa-Samuelson hypothesis are quite sensitive to alternative empirical estimation techniques (Chinn, 2000; Wang and Dunne, 2003; Tintin, 2009; Kakkar and Yan, 2012). In the studies investigating the Balassa-Samuelson hypothesis empirically, the long-run model is estimated using either the EG two-step single equation cointegration procedure or some form of generalized one-step error correction method (see Canzoneri et.al, 1999; Chinn, 2000; Lommatzsch and Tober, 2004; Bogoev et.al., 2008; Thomas and King, 2008; Tsen, 2011; Findreng, 2014). On the other hand, the maximum likelihood based rank test, proposed by Johansen (1988, 1991) and Johansen and Juselius (1990), is also popular amongst researchers. Though less widely used than the residual based cointegration tests, a number of studies on the

the Balassa-Samuelson effect use multivariate cointegration model to explore the plausible long-run association between productivity and real exchange rate (see Faruqee, 1995, Loko and Tuladhar, 2005; Konopczak and Torój, 2010; Jabeen et al., 2011 and Boreo et.al, 2015). Researchers have used both time-series empirical methods as well as pooled data estimation techniques to test the hypothesis (see MacDonald and Ricci, 2001; Drine and Rault; 2002, Lojshova; Baszkiewicz, 2004; Choudhari and Khan, 2005 Lee and Tang; 2007; Dumrongrittikul, 2012; Kakkar and Yan, 2013).

In the Balassa-Samuelson literature on Asia, there is a mix of evidence on the use of time-series or pooled data estimation methods. Bahmani-Oskooee and Nasir (2004) conduct an individual country-by-country study to test the Balassa-Samuelson hypothesis for a set of 44 countries (including six from East Asia). The study employs Bound testing approach for cointegration (ARDL model) for estimating the long-run effect of productivity on real exchange rate. The model, based on unrestricted (unconstrained) error correction test, is proposed by Pesaran and Shin (1999) and Pesaran, et al. (2001) to model the long-run and the short-run relationship between system variables simultaneously. Choudhari and Khan conduct a panel study on the Balassa-Samuelson hypothesis and examine two individual panels of low income and high income countries comprising of East Asian, African and Western Hemisphere states. Using Panel Dynamic Ordinary Least Squares (PDOLS) tests for cointegration, the study establishes long-run Balassa-Samuelson effect for two groups of countries by assuming homogenous cointegrating vectors within the group. Thomas and King (2008) investigate a set nine East Asian countries using a single equation residual based error correction model. The model is originally proposed by Zivot (1994). The existence (inexistence) of a valid cointegrating relationship between model variables is decided through an error correction process.

More recent studies from Asia use pooled data estimation techniques to investigate the Balassa-Samuelson hypothesis. Dumrongrittikul (2012) applies panel data estimation techniques to examine the BS hypothesis for a set 33 developing and developed countries. Using Group-Mean Panel Dynamic Ordinary Least Squares estimator suggested by Pedroni (2001b), the test allows for greater flexibility in the presence of heterogeneity of the co-integrating vectors since the panel includes countries from different regions. Ricci et al. (2013) analyse augmented version of the Balassa-Samuelson hypothesis for two individuals set of countries, advanced economies and newly industrialized emerging markets. Their sample includes 12 Asian countries. The estimation of equilibrium long-run cointegrating relationship between real exchange rate and its proposed determinants is undertaken using the panel dynamic ordinary least squares estimator developed by Stock and Watson (1993). The estimated relationship is further explored through panel error correction model to gauge the speed of convergence of real exchange rate towards its long-run equilibrium. Wang et al. (2016) base their study of the Balassa-Samuelson effect on a panel data set of 40 countries. Their sample includes 9 countries from Asia. The Balassa-Samuelson hypothesis is tested by employing an extended version of panel cointegration techniques, proposed by Bai and Carrion-i-Silvestre (2013), allowing for structural breaks and cross-sectional dependence. In their study, the long-run Balassa-Samuelson effect is estimated using group-mean panel cointegration estimator and the results suggest significant absence of the Balassa-Samuelson effect for the set of developing countries (only).

There are only a few studies on Asia that investigate the Balassa-Samuelson hypothesis by conducting alternative estimation methods. Interestingly, their findings yield quite conflicting results when the same log-run model is tested against two different empirical tests. Chinn (2000) observes valid existence of the Balassa-Samuelson effect, but of varying degree when the long-run model is tested though time series estimations methods and pooled data estimator. When conducting individual country analysis, the study employs single equation residual based error correction model of Phillips and Loretan (1991). Assuming that the model regressors are weakly exogenous, the author tests the Balassa-Samuelson model using non-linear least squares (NLS) regressions for nine East Asian states. The time-series analysis provides some evidence in support of the existence of the Balassa-Samuelson effect. For only three out of nine countries, the real exchange rate and productivity trend movements are found to comply with the theoretical predictions of the Balassa-Samuelson model. On the contrary, panel data estimation results are rather encouraging. Using panel NLS error correction model, the author confirms the valid existence of the Balassa-Samuelson effect for the panel countries. He finds significant mean-reversion in errors and convergence of real exchange rate towards its long-run equilibrium. Kakkar and Yan (2012) report considerable variation in their results when testing the Balassa-Samuelson hypothesis for six East Asian countries using alternative specifications of Kao and Pedroni panel cointegration tests. Though the two tests significantly favour the existence of the Balassa-Samuelson effect, the magnitude of effect is found to be highly varying under different circumstances. When the condition of homogenous cointegration vectors is imposed, the long-run Balassa-Samuelson coefficient is found to be 0.64 (approximately). Allowing for heterogeneous cointegrating vectors, the individual long-run coefficient estimates vary widely between 0.13 and 0.90.

My strategy of empirical verification of the Balassa-Samuelson hypothesis is more robust compared to the existing studies on Asia. To ensure better reliability of estimates, I use both single equation residual based cointegration procedures as well as multivariate cointegration method. I derive conclusions about the valid existence of the Balassa-Samuelson effect based on both the individual country-level and panel analysis. The details on testing procedures used in this dissertation are explained in chapter five.

4.2.4 Theoretical Foundations of the Model

(a) <u>Domestic Version of the Balassa-Samuelson Hypothesis</u>

While testing the Balassa-Samuelson hypothesis empirically, researchers have distinguished between the internal and external transmission mechanisms of the model. The internal mechanism of the hypothesis proposes that the sectoral productivity gap in a country may drive the internal relative prices of non-tradables. A faster productivity growth in the traded sector of the county may cause its relative price of non-tradables to rise. This effect is known in the literature as the Baumol-Bowen effect. Baumol and Bowen (1966) argued that within a country there is a broad tendency for the prices of service-intensive goods (education, health care, banking, etc.) to rise over time as, historically, productivity growth in these activities has tended to be slower than in more capital-intensive manufacturing industries. A sizeable number of non-Asian studies have tested the domestic version of the Balassa-Samuelson hypothesis (see Canzoneri et al., 1999; Egert, 2002, 2005; Lojschova, 2003; Mihaljek and Klau, 2004; Lee and Tang, 2007; Funda et al., 2008). These studies indicate that the domestic version of the model is the key driver of the standard the Balassa-Samuelson mechanism.

However, there is a serious dearth of studies investing the domestic version of the model for Asia. To my knowledge, the only credible work, conducting a multi-country analysis of Asia for the internal version of the Balassa-Samuelson hypothesis is done by Thomas and King (2008). The study investigates the long-run relationship between internal relative sectoral productivities and internal sectoral price ratios (augmented with a number of domestic demand side control variables like government consumption spending, GDP per capita, oil prices and terms of trade) for a set of nine East Asian countries and United States. In their study, seven out of ten countries reject the null hypothesis of no cointegration between model variables at ten percent or better significance level. However, when a general to specific approach is applied, i.e., clearly insignificant variables are dropped out of model, the parsimonious version of the model yields even more encouraging results. Nine out of ten countries reject the null of no cointegration, with Korea being the only exception.

The assumed value of relative labour intensities of nontraded sector $\left(\text{see the term } \gamma/_{\delta} \text{ in equation 2.9 of chapter two}\right)$ plays an important role in testing the domestic version of the Balassa-Samuelson model. If the non-traded sector is relatively more labour-intensive, i.e., $\gamma/_{\delta} > 1$, then even a balanced (proportionate) sectoral productivity growth may lead to an appreciation of relative prices of non-tradables (Froot and Rogoff, 1985). While testing the model empirically, Thomas and King (2008) assume an equi-proportionate relationship between biased sectoral productivity and relative sectoral prices, i.e., labour intensity is equal across sectors ($\gamma = \delta$ or $\gamma/_{\delta} = 1$) . However, the individual country estimates reveal that none of the subject economies meet this assumption. Almost always, the relative productivity bias in tradables appear to bear a disproportionate relationship with internal price ratio. This approach is consistent with the empirical findings of some earlier studies confirming the validity of a domestic Balassa-Samuelson effect but with disproportionate effects of productivity differentials on the internal real exchange rate (Mihaljik and Klau (2004); Egert, 2005; Lee and Tang, 2007; Funda et al., 2008).

(b) *Controlling for the Assumptions of the Model*

The Balassa-Samuelson hypothesis is widely criticized in recent empirical studies for its assumptions. Long-run PPP between inter-country traded sector prices[11] and inter-sectoral equalised wages[12] are two of the core assumptions of the model that are most widely tested in

[11] See Canzoneri et al., 1999; Egert, 2002b; Egert et al., 2003; Kovacs, 2003; Lojschova, 2003; Blaszkiewicz et al., 2004; MacDonald and Ricci, 2005; Lee and Tang, 2007; Garcia-Solanes et al., 2008.
[12] See Strauss and Ferris, 1996; Strauss, 1997, 1998; Nenovsky and Dimitrova, 2002; Lee, 2005.

literature and are found to be inexistent. In this section, I focus on the validity (invalidity) of the assumption that presumed long-run PPP between tradables prices across countries (see chapter eight for empirical verification of the assumption). The testing for this assumption is rather under-explored for Asia.

Few studies test the model by relaxing the PPP assumption and find quite different results. Ito.et al. (1999) investigate countries in Asia, Western Hemisphere and Oceania countries for the Balassa-Samuelson hypothesis and find sustained departures in traded sector prices vis-a-vis U.S. Through simple regressions, they find that ten out of eleven countries reportedly experience trend deviations in their relative prices of tradables against United States. The only exception was Korea that seems to be with the assumption of PPP. This finding casts shadow over one of the fundamental assumptions behind the Balassa-Samuelson hypothesis. Similarly, for a set of nine East Asian countries, Thomas and King (2008) investigate the equi-proportionate relationship between home tradables prices and corresponding U.S. prices as proposed by the Balassa-Samuelson hypothesis. Using single equation error correction models, their findings reveal a valid long-run co-movement amongst model variables, but only for half of the sample countries. Kakkar and Yan (2012) examine the assumption of tradable sector PPP for East Asia and find mixed evidence. Their empirical analysis provide some support for PPP between inter-country traded sector prices when the restriction of homogeneity on cointegration vectors is imposed, but generally weaker under heterogeneous vectors.

The existing empirical evidence on this issue is too limited to develop a clear understating on them. The sensitivity of long-run real exchange rates towards internally biased relative productivity of tradables and trend departures of inter-country tradables prices from PPP equilibrium are insufficiently examined for Asia. In chapters seven and eight of the thesis, I extend my empirical research in two areas by pursuing my line of inquiry on domestic version of the Balassa-Samuelson hypothesis and modified version of the standard (international) Balassa-Samuelson model (by relaxing the assumption of PPP in tradables) respectively. Thus, I try to seek empirical evidence on the Balassa-Samuelson hypothesis but with altered theoretical specifications. The consistency of estimates is then evaluated against alternative empirical models. The contribution is valuable as the assumption of PPP in tradables has so far been tested empirically for Asia by barely a handful of studies.

Below, TABLE 4.1 provides a summary of the existing literature examining the Balassa-Samuelson hypothesis for East Asia and South Asia. The table provides readers with a quick view of each study highlighting the sample countries, scheme of sectoral division, measures of real exchange rate, verification of underlying model assumptions and final conclusion about the existence of the Balassa-Samuelson effect.

TABLE 4.1: Summary of Critical Features of Noticeable BS Studies on Asia

Critical Features of the Study	Bahmani-Oskooee and Rhee (1996)	Ito et al. (1999)	Chinn (2000)
(a) Sample Countries	Korea	18 APEC countries (11 from Asia, 4 from Western Hemisphere and 3 from Oceania)	China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand.
(b) Sample Data Set	Quarterly 1979-93	Annual 1973-93. Start and end dates vary from country to country.	Annual 1970-92. Start date varies from country to country.
(c) Scheme of Sectoral Division	-	Highly aggregated, T = Manufacturing, NT = Services	T = LDCs: Manufacturing, US & Japan: Industry, mining, transportation & agriculture NT = LDCs: Services, construction, mining and transportation, US & Japan: Services, construction, government.
(d) Measure(s) of RER	GNP deflator based RER.	GDP deflator based RER.	GDP deflator based RER.
(e) Domestic or/and International Version of the Model	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.
(f) Controlling for the Idealist Assumptions of the Model	No	The countries' exchange rate are allowed to deviate from PPP.	Demand-side factors are allowed to explain real exchange rate movements.
(g) Empirical Estimation Methodology	Johansen-Juselius Maximum-Likelihood Cointegration Test	Simple OLS regression model	Time-series and panel error correction regressions.
(h) Conclusion	Valid existence of BS effect is found	Mixed support is found in favour of BS hypothesis	Valid existence of BS effect is confirmed, more robustly through panel data estimation methods.

Critical Features of the Study	Drine and Rault (2002)	Wang Dunne (2003)	Bahmani-Oskooee and Nasir (2004)
(a) Sample Countries	India, Indonesia, Korea, Philippines, Singapore and Thailand	Indonesia, Japan, Korea, Malaysia, Philippines, Singapore and Thailand.	A set of 44 developed, developing and less-developed countries including six countries from East and South Asia.
(b) Sample Study Period	Annual 1983-98	Quarterly 1973-96. Start date varies for some countries.	Annual 1960-90
(c) Scheme of Sectoral Division	T = Manufacturing sector and agriculture, hunting, forestry and fishing. NT = Services	Inter-country tradable sector productivity gap is taken into account only. The tradable sector is represented by manufacturing sector of each country and for those countries where manufacturing output data is not available, real GDP is used as an alternative measure.	-
(d) Measure(s) of RER	CPI based effective RER	CPI based bilateral RER.	GDP deflator-based RER
(e) Domestic or/and International Version of the Model	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.
(f) Controlling for the Idealist Assumptions of the Model	The countries' exchange rate are allowed to deviate from PPP.	No	No
(g) Empirical Estimation Methodology	Multivariate (Johansen) and panel (Pedroni) cointegration tests.	Johansen-Juselius ML cointegration method and generalized variance decomposition test.	Autoregressive Distributed Lag (ARDL) bound testing cointegration test and Error Correction model.
(h) Conclusion	BS effect is invalidated for the sample countries.	Except Singapore, very little evidence is obtained in support of BS effect for other sample countries.	For Asia 4 out of 6 countries display the valid existence of BS effect.

Critical Features of the Study	Choudhri and Khan (2005)	Thomas and King (2008)	Olson (2009)
(a) Sample Countries	The study includes 14 Asian, African and South American countries at high, low- and medium-income levels.	China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand.	China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan, and Thailand.
(b) Sample Study Period	Annual 1976-94.	Annual 1960-2004. Start date varies from country to country.	Annual 1970-2007
(c) Scheme of Sectoral Division	T = Manufacturing and agriculture NT = All other sectors	T = Manufacturing NT = Services, construction and utilities.	-
(d) Measure(s) of RER	CPI based RER	GDP deflator based RER	CPI based RER
(e) Domestic or/and International Version of the Model	International version of the hypothesis is tested only.	Both domestic and international versions of the hypothesis are tested.	International version of the hypothesis is tested only.
(f) Controlling for the Idealist Assumptions of the Model	No	i. Long run co-movement between cross-country tradables prices (PPP) is investigated. ii. Non-tradable price component in tradable prices is taken into account.	No
(g) Empirical Estimation Methodology	Two residual-based Pedroni tests and Panel DOLS estimator.	Time-series error correction regressions	Single-equation cointegration model and impulse response function.
(h) Conclusion	The results strongly suggest that the BS hypothesis is in operation.	Authors obtain significantly stronger support in favour of valid existence of BS effect.	The model augmented with demand-side shocks supports the valid existence of BS effect for subject countries.

Critical Features of the Study	Tsen (2011)	Dumrongrittikul (2012)	Chowdhury (2012)
(a) Sample Countries	Japan, Korea, and Hong Kong	33 countries in total. 17 developing countries and 16 developed countries. Fourteen countries in sample are from Asia.	Seven low-income SAARC economies (Bangladesh, Bhutan, Nepal, India, Maldives, Pakistan and Sri Lanka)
(b) Sample Study Period	Quarterly 1960-2009. Start date varies from country to country.	Annual 1970-2008	Annual 1950-2007. Start date varies from country to country.
(c) Scheme of Sectoral Division	-	Each country's real economy is divided into seven distinct sectors. T and NT sectors vary from country to country.	-
(d) Measure(s) of RER	CPI based bilateral RER.	CPI based RER	GDP Deflator-based RER
(e) Domestic or/and International Version of the Model	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.
(f) Controlling for the Idealist Assumptions of the Model	No	No	No
(g) Empirical Estimation Methodology	Engle-Granger residual based test for cointegration, Johansen (1988) cointegration method and the SL cointegration method and the generalised forecast error variance decomposition.	Four residual-based Pedroni tests and Group-Mean Panel DOLS and time-series DOLS estimators.	Autoregressive Distributed Lag (ARDL) bound testing cointegration approach and Error Correction model.
(h) Conclusion	Considerable amount of support is yielded for hypothesis.	For the set of developing countries (only), strong evidence in support of the BS effect is found.	Dominant inexistence of BS effect is found.

Critical Features of the Study	Kakkar and Yan (2012)	Ricci et al. (2013)	
(a) Sample Period	Hong Kong, Indonesia, Korea, Malaysia, Singapore, and Thailand.	The sample set comprises of a mix of industrial economies and emerging market. 11 countries are taken from Asia.	
(b) Sample Study Period	Annual 1980-2001	Annual 1980-2004	
(c) Scheme of Sectoral Division	T = manufacturing; mining & quarrying; ocean and air transport; wholesale & retail trade; and financing, insurance & business services. NT = utilities; construction; real estate; community, social, and personal services; land transport and communication; and restaurants	T = agriculture, hunting, forestry, and fishing; mining, manufacturing, and utilities; and transport, storage, and communication. NT = construction; wholesale and retail trade; and other services.	
(d) Measure(s) of RER	CPI based RER.	CPI based effective RER.	
(e) Domestic or/and International Version of the Model	International version of the hypothesis is tested only.	International version of the hypothesis is tested only.	
(f) Controlling for the Idealist Assumptions of the Model	Long run co-movement between cross-country tradable prices (PPP) is investigated.	Demand-side determinants are also modelled against RER.	
(g) Empirical Estimation Methodology	Kao and Pedroni cointegration tests for detecting plausible long run association. The long run elasticities are calculated through Panel DOLS estimator.	Panel DOLS and Panel Error Correction model.	
(h) Conclusion	Valid BS effect is concluded.	Valid BS effect is concluded but with small magnitude (particularly for industrialized countries).	

CHAPTER 5: INVESTIGATING THE STANDARD VERSION OF THE BALASSA-SAMUELSON HYPOTHESIS

5.1 Introduction

This study explores the productivity-based, long-run determinants of real exchange rates. As this chapter is the initial empirical investigation of the productivity-real exchange rate nexus, I will test their long-run association under the most basic and standard theoretical specification of the Balassa-Samuelson (BS) model. However, consistency and reliability of results is not compromised. Special care is taken to ensure the robustness of results, using alternative measures of real exchange rate and various econometric methods, whilst testing the hypothesis empirically.

Purchasing Power Parity (PPP) theory predicts that relative prices in two countries determine the nominal exchange rate in the long-run. According to this theory, the real exchange rate in long-run should be equal to one. Any deviation of the real exchange rate from unity should be transient. An important modification to PPP theory is the BS hypothesis. The BS hypothesis predicts that permanent deviations of real exchange rates from unity can occur due to productivity differences in the traded and nontraded sectors between two countries. Even if PPP holds for tradables prices, nontraded sector price differences arise (due to imbalanced sectoral productivity growth patterns), thus driving long-run real exchange rate movements.

The role of exchange rate in the monetary policy framework for emerging Asian economies is not a new issue, but it remains a hot topic for academic research. Over the last twenty years, East and South Asian economies have been faced with serious real exchange rate appreciation against major international currencies, particularly the U.S. dollar. With this in mind, my study will investigate real exchange rate behavior for each of the following member states of ASEAN+3 and SAARC territories.

TABLE 5.1: List of ASEAN and SAARC Countries

ASEAN	North East Asian Dialogue Partners of ASEAN	SAARC
Indonesia Malaysia Philippines Thailand Singapore	Hong Kong Japan Korea	Pakistan Sri Lanka

Depending upon the availability of data, my empirical strategy is to take each of the ten countries identified in TABLE 5.1 and estimate the relationship between real exchange rate and productivity, using the U.S. as the reference trading partner. Selection of the U.S. as the reference economy is supported by a number of studies investigating emerging Asian countries for BS theory (e.g., Ito et al., 1999; Chinn; 2000; Choudhri and Khan, 2005; Thomas and King, 2008; and Dumrongrittikul, 2012).

The remainder of the chapter is devoted to (a) econometric procedures employed to verify the existence of the BS effect, (b) salient features of model variables and sample data sets, and (c) individual country analysis for eight Asian economies verifying the BS hypothesis (Singapore and Hong Kong are excluded for this chapter). The first country that I examine is Korea. I will do this in great detail so that the reader will understand the many steps that are involved in testing the BS hypothesis. Once I have established the protocol, I will then proceed to the other countries, but will provide only a summary report of my empirical findings.

5.2 Empirical Framework for Verifying the Balassa-Samuelson Hypothesis

In its simplest form, the BS hypothesis posits the following empirical relationship to estimate long-run (permanent) departure of real exchange rate from its PPP based equilibrium, driven by imbalanced sectoral productivity growth patterns of a home economy (Asia) against a foreign economy (U.S.).

(5.1) $rer_t = \alpha + \beta \tilde{a}_t + \varepsilon_t$,

where $rer_t = (e_t + p_t^*) - p_t =$ bilateral real exchange rate of home against U.S. (represented by *) at time t is defined as the product of their bilateral nominal exchange rate and relative aggregate prices in national currencies.

$\tilde{a}_t = -[(a_t^T - a_t^{NT}) - (a_t^{T*} - a_t^{NT*})] =$ cross-country relative sectoral productivity growth differential at time t.

The small letters are representing the variable series in logarithmic form.

Equation (5.1) is my starting point for investigating $\tilde{a}$ as a potential long-run determinant of rer. However, the estimation of the relationship between rer and $\tilde{a}$ is complicated by the time series properties of these variables. I focus on the following case to determine if the standard version of the BS hypothesis as stated above can explain the observed behavior of rer.

5.2.1 Real Exchange Rate and Productivity are both Non-Stationary but Cointegrated

When rer and $\tilde{a}$ are non-stationary in levels and are co-integrated, their relationship can be modelled using a cointegration approach.

Non-stationary (trended) time-series data can be a major problem when estimating relationships between time series variables. It is well known that trends, either stochastic or deterministic, may cause spurious regressions, yielding unreliable estimates. However, most macroeconomic time-series are subject to some type of trend. A real breakthrough in time-series econometrics came with the concept of "cointegration" in the early 1980s. The concept was first introduced by Granger (1981). Afterwards, Engle and Granger (1987), in their seminal paper, provided a firm theoretical base for modelling and estimating cointegrated non-stationary time-series. Cointegration analysis allows non-stationary data to be used in a meaningful manner, so that spurious results are avoided. The method provides applied econometrics an effective formal framework for testing and estimating long-run models from actual time-series data.

The estimation and testing of a co-integrating relationship can be performed in a single-equation framework (e.g., Engle-Granger type residual-based framework) or in a multivariate (VAR-based) framework (e.g., Johansen-Juselius Maximum Likelihood procedure). The two types of cointegration models have their own strengths. Let us take a generic overview of how the two cointegration approaches differ in their basic mechanism whilst establishing long-run cointegrating relationships between two or more time series.

But before looking into the cointegration tests formally, let's briefly discuss the two unit root tests I employ to determine the order of integration of rer and $\tilde{a}$. In general, a linear combination of two I(1) time series will have an order of integration equal to one (indicating that there is no long-run relationship between the two), or zero. In the latter case, the series are said to be "cointegrated" and we can use econometric procedures to estimate the long-run relationship. However, the first step in determining whether variables are cointegrated is to test for their order of cointegration. It is standard practice to determine the order of integration of variables before subjecting them to formal cointegration tests.

(i) STEP 0-Test for Unit Roots

It is well known that different tests can produce different results. For this reason, it is a wise idea to see how consistent my results are across two different unit root test procedures.

(a) Augmented Dickey-Fuller Unit Root Test (ADF) test

The general form of the Augmented Dickey-Fuller (ADF) test is:

(5.2) $$\Delta y_t = c + \delta t + \phi y_{t-1} + \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \varepsilon_t \,,$$

where c and t are deterministic regressors that allow for either a constant term or a time trend if the series is trend stationary, or a drift term and a quadratic time trend if the series is difference stationary. In addition, lagged differences are added to control for the effects of serial correlation, which can otherwise invalidate hypothesis testing.

The two series will be tested individually against the null hypothesis of unit root, i.e., $H_0: \phi = 0$ versus the alternative hypothesis of $H_1: \phi < 0$, i.e., the series is generated through a stationary process.

(b) Dickey-Fuller Generalized Least Squares (DF-GLS)[13] Unit Root Test

DF-GLS unit root test is essentially the ADF unit root test, except that the time series (rer or $\tilde{a}$) is transformed via a generalized least squares regression before performing the unit root test. Using equation (5.2), the DF-GLS test is performed analogously but on GLS-detrended data. The two series are tested individually against the null hypothesis of unit root, i.e., $H_0: \phi = 0$ versus the alternative hypothesis of $H_1: \phi < 0$.

[13] DF-GLS has two specifications. In GLS de-trending the subject series is regressed on a constant and a linear trend and the resultant residual series is used in a standard DF regression. On the other hand, for GLS demeaning, the series is regressed on a drift term only and the resultant residuals series will be used in standard DF regression.

Whilst conducting the country studies in this chapter as well as for forthcoming chapters, it is expected that I may obtain contrasting results from ADF and DF-GLS unit root tests, leaving me indecisive as to the true order of integration of the rer and $\tilde{a}$ series. However, I cannot overlook the problem of size distortion, inherent to conventional unit root tests, thus making these tests to produce misleading results sometimes (c.f., Reed, 2016). Running 10,000 simulations of sample size 100, the author reveals that ADF, DF-GLS and Phillips-Perron tests are commonly oversized. Reed and Smith (2016) demonstrate that the reason for this is the cointegrated series can be represented by univariate time series representation with an MA term. As is well known (Ng and Perron, 2001), the presence of an MA term causes unit root tests to over-reject the null hypothesis of non-stationarity. Adding sufficient lags to the Dickey-Fuller testing specification fixes the problem. However, standard approaches to selecting the number of lags (e.g., using information criteria such as the AIC, SIC, or Modified AIC; or testing for serial correlation) results in too few lags being included in the testing specification. The result is that variables that are cointegrated may in fact produce test results that indicate that one or both of them are stationary. As a result, while I will test for unit roots in the individual series, my determination of whether a series are cointegrated will depend on direct tests for cointegration.

(ii) Single Equation Cointegration Approach

A number of different methods for estimating cointegration are suggested in the literature. Among these, the class of residual-based cointegration tests is the most popular, being simple in computation with straightforward interpretation in terms of economic theory. Under the residual-based cointegration approach, the existence of cointegration between two macroeconomic variables implies that they have a meaningful association. It is this association which prevents the residuals from becoming larger and larger in the long-run.

From the group of residual based cointegration tests, the Engle-Granger (EG) type static long-run regression models have received immense popularity, formally introduced by Engle and Granger (1987). The EG type long-run ordinary-least-squares (OLS) residual based estimator is acknowledged as both consistent and highly efficient (see Stock, 1987). Following the unit-root testing approach, the literature has proposed tests for the null hypothesis of no cointegration (e.g., Engle and Granger, 1987; Phillips and Ouliaris, 1990) as well as tests for the null hypothesis of cointegration (e.g., Hansen, 1992; Shin, 1994).

Another type of test is based on an Error Correction (EC) representation. Error Correction Models (ECMs) are based on the behavioural assumption that two or more time series exhibit an equilibrium relationship that determines both short- and long-run behaviour. The transient disequilibrium is captured by the inclusion of a disequilibrium term in the ECM which picks up the extent of departure of the model time series from their long-run equilibrium. When that happens, and assuming the variables are cointegrated, one or both of the variables will adjust to restore the equilibrium.

Previous studies of the BS hypothesis have used either the EG two-step single equation cointegration test or some form of generalized one-step error correction method (see Canzoneri et al., 1999; Chinn, 2000; Lommatzsch and Tober, 2004; Bogoev et al., 2008; Thomas and King, 2008; and Findreng, 2014).

Keeping in view the amount of popularity enjoyed by above discussed single equation cointegration tests, I will test the BS theory using both approaches. I will conduct the single equation cointegration analysis in two distinct steps, though inter-connected but acting independently. Each step will be confirming/invalidating the perceived BS effect in its own unique way.

(a) STEP 1-Test for Cointegration

STEP 1 of the single equation cointegration approach consists of verifying the existence/inexistence of a valid cointegrating relationship as suggested by equation (5.1). It consists of two individual tests: (i) an EG residual-based cointegration test, testing the residuals from equation (5.1) for being mean reverting; and (ii) an error correction representation, modelling errors from equation (5.1) in an autoregressive function of first-differenced rer and $\tilde{\alpha}$. The two tests do not allow for reverse causality or endogeneity in model variables. Thus, it is assumed that $\tilde{\alpha}$ is causing rer but not vice versa. Let's briefly elaborate the significance of the two tests[14] in the form of the conditions necessary to establish cointegration between rer and $\tilde{\alpha}$.

S1.A. *EG Residual-Based Cointegration Test:* As the first formal step of cointegration, I will test the residuals from a static cointegration regression of equation (5.1). The testing is done through EViews, using an Engle Granger single cointegration test specification. If the residuals exhibit short memory, displaying significant mean reversion, the rer and $\tilde{\alpha}$ series will be concluded as cointegrated in the long-run.

S1.B. *Estimating Error Correction Model (ECM):* Here I will model the residuals (obtained from equation 5.1) in an autoregressive function of first-differenced rer and $\tilde{\alpha}$, called the Error Correction Model (ECM). The association between cointegration and the ECM stems from the Granger representation theorem. The theorem suggests that a statistically significant error correction representation is likely to hold between two or more integrated time series if they are cointegrated. This implies that statistically significant estimates of the error correction term can also be taken in support of a valid cointegration between rer and $\tilde{\alpha}$. The residual series from equation (5.1) will serve as a measure of disequilibrium which captures and reveals the extent of

[14] Each test under single equation cointegration approach is abbreviated with 'S' symbolizing single equation test.

short-lived deviations in rer from the long-run equilibrium. Obtaining a negative and statistically significant Error Correction (EC) coefficient implies adjustments are made by the model time series at some significant (statistically) speed, so that part of the rer convergence to its equilibrium levels is achieved by these adjustments.

(b) STEP 2-Estimating the Long-Run Relationship (BS Effect)

Having established that rer and $\tilde{a}$ are cointegrated, the second and final step of the single equation cointegration procedure consists of estimating whether the long-run relationship between rer and $\tilde{a}$ is consistent with the theoretical prediction of the BS model. This is indicated by a positive and statistically significant sign on the estimated coefficient of the $\tilde{a}$ variable. The long-run BS coefficient will be estimated using the following two estimators.

S2.A. *Estimating Long-Run BS Coefficient through FMOLS:* Fully Modified Ordinary Least Squares (FMOLS) cointegration regression estimator is used to obtain the point estimates for the BS effect.

S2.B. *Estimating Long-Run BS Coefficient through DOLS*: To check the robustness of estimates obtained through FMOLS, the long-run BS coefficient is also estimated through the Dynamic Ordinary Least Squares (DOLS) cointegration regression estimator.

(iii) Multivariate Cointegration Approach

In addition to the single equation cointegration approach for estimating long-run equilibrium, Johansen (1988, 1991) and Johansen and Juselius (1990) provides a Maximum Likelihood (ML) based, system of equations approach. The main advantage of the Johansen ML method is that it allows one to detect situations where more than one variable adjusts to restore long-run equilibrium. In other words, it does not impose the condition of weak exogeneity. This characteristic of the Johansen ML approach stands in contrast to the single equation cointegration approach, which assumes that $\tilde{a}_t$ does not change in response to departures from long-run equilibrium with rer_t.

Though less widely used than residual based single equation cointegration tests, a number of studies of the BS hypothesis use the multivariate cointegration model to explore the long-run association between $\tilde{a}$ and rer (e.g., Faruqee, 1995, Loko and Tuladhar, 2005; Konopczak and Torój, 2010; Jabeen et al., 2011; and Boreo et al., 2015).

The Johansen ML test can be seen as a generalization of the Augmented Dickey-Fuller (ADF) unit root test under a multivariate system of equations. This generalization allows one to investigate the linear combination of variables for a unit root. In the event of more than two model variables, the maximum likelihood estimation approach of Johansen method can identify all possible cointegrating vectors. For instance, if there are *n* variables that all have unit roots, there can be at most *n-1* valid cointegrating vectors. The existence of valid cointegrating vector(s) is evaluated using two tests: the Trace statistic test and the Maximum Eigenvalue test. For a given value of $0 \leq r^* < n$, the Trace test tests the null hypothesis that the number of cointegrating vectors is less than or equal to r^*, against the alternative, that the number of cointegrating vectors is greater than r^*. The Maximum Eigenvalue test, on the other hand, tests the null hypothesis of r^* cointegrating vectors against the alternative hypothesis of $r^* + 1$ cointegrating vectors. Asymptotic critical values can be found in Johansen and Juselius (1990). For my analysis, I will run the test using EViews.

As I already have stated, unlike the single equation cointegration approach, the multivariate cointegration test does not rule out the possibility of endogeneity in model variables. Being multivariate in nature, the Johansen ML test allows for reverse causality between model variables; i.e., that changes in $\tilde{a}$ cause changes in rer. This provides another reason to estimate a

Vector Error Correction (VEC) model, which I elaborate in STEP 2 below. Here are the steps involved in estimating cointegration between rer and $\tilde{\alpha}$ in the multivariate framework.

(a) *STEP 1-Test for Cointegration*

Similar to the single equation cointegration approach, the first step using the multivariate cointegration method involves detecting whether there exists a valid cointegrating relationship between rer and $\tilde{\alpha}$ using a maximum likelihood based rank test.

M1. *Maximum Likelihood Based Rank Test:* At first place, I will determine the rank of the Johansen test, i.e., how many valid cointegrating relationships/vectors are established by the model. For this chapter, keeping in view the two model variables (rer and $\tilde{\alpha}$), I will conclude that the two variables are cointegrated if and only if there is one cointegrating vector (as opposed to zero or two).

(b) *STEP 2-Test for Exogeneity through VEC Model*

The next step is to establish a VAR-based error correction representation. Unlike the single equation ECM, I can use a Vector Error Correction Model (VECM) to determine whether $\tilde{\alpha}$ is exogenous (weakly) with respect to rer or not. VEC test models inter-relationship of two unit root time series in a VAR specification. Analogous to single equation ECM, the VEC model is a system of autoregressive equations, fitted to the first-differences of the non-stationary time series (rer and $\tilde{\alpha}$) together with lagged residuals from equation (5.1), representing deviations from long-run equilibrium.

Verifying the exogeneity of $\tilde{\alpha}$ with respect to rer involves estimating two error correction models individually, in which the change in rer and $\tilde{\alpha}$, respectively, are alternatively employed as regressands. This will be done automatically under VEC model, estimated through EViews. To validate the condition of weak exogeneity, the following two conditions are required to hold.

M2.A. *EC coefficient in the rer equation (EC_{rer}) is negative and statistically significant:* If the valid cointegrating vector (under the Johansen ML rank test) is generated through $\tilde{a}$ causing rer, this requires a negative and statistically significant EC coefficient (EC_{rer}) in the rer equation.

M2.B. *EC coefficient in $\tilde{a}$ equation ($EC_{\tilde{a}}$) is statistically insignificant:* To confirm the one-way causality between rer and $\tilde{a}$ -- i.e., that $\tilde{a}$ is causing rer and not vice versa -- one needs to obtain a statistically insignificant EC coefficient ($EC_{\tilde{a}}$) in the $\tilde{a}$ equation.

(c) <u>STEP 3-Estimating the Long-Run Relationship (BS Effect)</u>

Similar to the single equation cointegration approach, as a final step, the multivariate cointegration approach also estimates the long-run relationship (BS effect) between rer and $\tilde{a}$.

M3. *A Positive and Statistically Significant Long-Run BS coefficient:* The VEC model when estimated through EViews yields a long-run point estimate of $\tilde{a}$. A positive and statistically significant coefficient on $\tilde{a}$ in the cointegrating equation confirms the valid existence of the BS effect.

TABLE 5.2 summarizes the conditions above necessary to establish support for the BS hypothesis in the individual country studies. It also provides analogous tests and conditions when the individual country data are pooled (cf. Section 5.6 of this chapter.).

TABLE 5.2: Conditions for Establishing Balassa-Samuelson Effect

The BS effect is investigated empirically using both individual country data and pooled data. The two approaches employ single equation cointegration method as well as system of equations approach. The conditions necessary for establishing valid BS effect under both types of empirical frameworks are outlined below.

I-Time-Series Estimation Methods (Individual Country Study)

(a) STEP 0:Tests for Unit Root in Individual Series

The two time series rer and $\tilde{a}$ will be tested individually for their order of integration using the following two unit root tests.

(i) Augmented Dickey-Fuller Unit Root Test(ADF) test

The general form of ADF unit root test is:

(5.2)' $\Delta y_t = c + \delta t + \phi y_{t-1} + \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \varepsilon_t$, $y = rer$ and $\tilde{a}$

The two series will be tested individually against the null hypothesis of unit root, i.e., $H_0\!: \phi = 0$ versus the alternative hypothesis of $H_1\!: \phi < 0$; i.e., the series are generated through a stationary process.

(ii) Dickey-Fuller Generalized Least Squares (DF-GLS)Unit Root Test

Using equation (5.2)', the DF-GLS test is performed analogously but on GLS-detrended data. The two series will be tested individually against the null hypothesis of unit root, i.e., $H_0\!: \phi = 0$ versus the alternative hypothesis of $H_1\!: \phi < 0$.

(b) Single Equation Cointegration Approach

STEP 1:Test for Cointegration

S1.A. *Residual-Based Single Equation Test:* For establishing long-run association between rer and $\tilde{a}$, the residuals obtained through regressing rer on $\tilde{a}$ (equation 5.1) must be generated through a stationary process.

(5.3) $\quad \epsilon_t = rer_t - \hat{\alpha} - \hat{\beta}\,\tilde{a}_t$,

ϵ_t represents the residual series obtained from regressing rer on $\tilde{a}$ using OLS.

(5.4) $\quad \Delta\epsilon_t = \delta + \varphi\epsilon_{t-1} + \sum_{j=1}^{p}\omega_j\,\Delta\epsilon_{t-j} + u_t$,

Under the condition of mean reverting residuals, the estimated statistics for φ in equation (5.4) will be compared against the test critical values at the 10% significance level. The associated critical values for the *t*-statistic for $\hat{\varphi}$ are taken from MacKinnon (1996).

S1.B. *Estimating the Error Correction Model (ECM):* A statistically significant error adjustment (correction) parameter is analogous to establishing a valid long-run association between rer on $\tilde{a}$. A negative and statistically significant EC coefficient implies significant speed of adjustment of model variables, so that short-lived fluctuations of rer could be corrected and rer may converge to its long-run equilibrium.

(5.5) $\quad \Delta rer_t = \gamma + \rho(rer_{t-1} - \alpha - \beta\,\tilde{a}_{t-1}) + \sum_{i=1}^{k}\mu_i\,\Delta rer_{t-k} + \sum_{i=1}^{k}\lambda_i\,\Delta\,\tilde{a}_{t-k} + \nu_t$,

$EC\ Coefficient = \rho,\ \rho < 0$ and is statistically significant.

<u>*STEP 2: Estimating the Long-Run Relationship (Balassa-Samuelson Effect)*</u>

S2.A. A positive and statistically significant BS coefficient, β, is estimated using Fully Modified OLS (FMOLS).

S2.B. A positive and statistically significant BS coefficient, β, is estimated using Dynamic OLS (DOLS).

(c) Multivariate Cointegration Approach

<u>*STEP 1:Test for Cointegration*</u>

Maximum Likelihood Based Rank Test: The condition for establishing a cointegrating relationship between rer and $\tilde{a}$ comes from Johansen (1988, 1991) and Johansen-Juselius (1990)'s ML cointegration procedure. As this is a system of equations, it allows the possibility of more than one cointegrating vector.

(5.6) $\Delta X_t = \mu + \Pi\ X_{t-1} + \Gamma_1\,\Delta X_{t-1} + \cdots\cdots\cdots + \Gamma_{p-1}\,\Delta X_{t-p+1} + \varepsilon_t$,

where X is a 2×1 vector of rer and $\tilde{a}$.

Cointegration necessitates that the rank of the matrix Π , representing the number of cointegrating vectors, must meet the following condition:

(5.7) $rank(\Pi) = 0 < r < n$,

where n is the number of model variables. If there are n variables and there are n cointegrating vectors, then the model time-series are stationary.

The existence of valid cointegration is detected using the Trace and Maximum Eigenvalue statistics. Given $0 \le r^* < n$, the Trace test tests the null hypothesis of no more than r^* cointegrating vectors against the alternative hypothesis of more than r^* cointegrating vectors. The Maximum Eigenvalue test, on the other hand, tests the null hypothesis of r^* cointegrating vectors against the alternative hypothesis of $r^* + 1$ vectors.

M1.A *The Trace test indicates that there is one cointegrating vector.*

M1.B *The Max Eigenvalue test indicates that there is one cointegrating vector.*

STEP 2: Test for Exogeneity through VEC Model

Existence of a valid BS effect also requires $\tilde{a}$ to be weakly exogenous. For this purpose, the Vector Error Correction (VEC) model will be estimated through two equations. The following conditions are required to prove the exogeneity (weak) of $\tilde{a}$.

M2.A. *EC coefficient in rer equation (EC_{rer}) is negative and statistically significant*

(5.8-a) $\Delta rer_t = \gamma_1 + \theta_{rer}(rer_{t-1} - \alpha - \beta\ \tilde{a}_{t-1}) + \sum_{i=1}^{k} \mu_{1i}\,\Delta rer_{t-k} + \sum_{i=1}^{k} \lambda_{1i}\,\Delta\ \tilde{a}_{t-k} + \nu_{1t}$,

$EC_{rer} = \theta_{rer}$, $\theta_{rer} < 0$ and is statistically significant. This implies that rer adjusts to deviations from long-run equilibrium in rer and $\tilde{a}$.

M2.B. *EC coefficient in $\tilde{a}$ equation ($EC_{\tilde{a}}$) is statistically insignificant*

(5.8-b) $\Delta\,\tilde{a}_t = \gamma_2 + \theta_{\tilde{a}}(rer_{t-1} - \alpha - \beta\,\tilde{a}_{t-1}) + \sum_{i=1}^{k}\lambda_{2i}\,\Delta rer_{t-k} + \sum_{i=1}^{k}\mu_{2i}\,\Delta\,\tilde{a}_{t-k} + \nu_{2t}$,

$EC_{\tilde{a}} = \theta_{\tilde{a}}$ *is statistically insignificant*. This implies that $\tilde{a}$ is weakly exogenous, i.e., $\tilde{a}$ is unaffected by deviations from long-run equilibrium in rer and $\tilde{a}$.

STEP 3: Estimating the Long-Run Balassa-Samuelson Coefficient

M3. A positive and statistically significant BS coefficient in error correction term under VEC model will validate the BS effect.

Evidence in Favor of BS Hypothesis under Time-Series Estimation Approach

Single Equation Cointegration Approach	*Multivariate Cointegration Approach*
"YES" (i) At least one of S1.A and S1.B; and (ii) Both S2.A and S2.B hold valid. "MIXED" (i) At least one of S1.A and S1.B; and (ii) S2.A or S2.B (but not both) hold valid. If any situation other than 'YES' and 'MIXED' occurs, it will be concluded as 'NO'.	"YES" (i) At least one of M1.A and M1.B; and (ii) Both M2.A and M2.B; and (iii) M.3 hold valid. "AMBIGUOUS" (i) At least one of M1.A and M1.B; and (ii) M2.A; and (iii) M.3 hold valid. If any situation other than 'YES' and 'AMBIGUOUS' occurs, it will be concluded as 'NO'.

II-Pooled Data Estimation Methods

STEP 0: Test for Unit Root in Individual Panels

The individual panels in pooled data will be tested for their order of integration using the following two unit root tests and one stationarity test.

(i) Combining p-values Test- Fisher-Type Panel Unit Root Tests
- Fisher-ADF Unit Root Test
- Fisher-PP Unit Root Test

(ii) Hadri LM Residual-Based Stationarity Test

STEP 1:Test for Cointegration

(a) Residual Based Cointegration Test

<u>*Pedroni Residual-Based Single Equation Cointegration Test*</u>: Under single equation cointegration approach, I employ the following seven Pedroni (1999, 2004) residuals-based tests to decided cointegration between model variables for the panel of my subject Asian countries.

1. Panel v statistic
2. Panel ρ statistic
3. Panel PP statistic
4. Panel ADF test
5. Group ρ statistic
6. Group PP statistic
7. Group ADF test

P1.A. The guideline for establishing a valid cointegration between rer and $\tilde{a}$ is if a majority of the above stated seven tests reject the null hypothesis of no cointegration with appropriate statistical significance. This will be taken as an evidence in support of long-run co-movement between rer and $\tilde{a}$.

(b) Maximum Likelihood-Based Multivariate Cointegration Approach

<u>*Fisher-Johansen Combined Multivariate Cointegration Test*</u>*:* Similar to Johansen-Juselius (1990) ML cointegration test, Fisher-Johansen Combined cointegration method is also a maximum likelihood based cointegration test. The evidence on cointegration is yielded through the rank of the model, tested against the following two tests:

P1.B: Johansen-Fisher Trace Test
P1.C: Johansen-Fisher Max Eigenvalue Test

STEP 2. Estimating the Long-Run Balassa-Samuelson Coefficient

P2 (A). A positive and statistically significant BS coefficient, β, established through the Panel Fully Modified OLS (PFMOLS) estimator.

P2 (B). A positive and statistically significant BS coefficient, β, established through the Panel Dynamic OLS (PDOLS) estimator
Evidence in Favor of BS Hypothesis under Pooled Data Estimation Approach
"YES" (i) At least two of P1.A, P1.B and P1.C; and (ii) Both P2.A and P2.B hold valid. "MIXED" (i) At least two of P1.A, P1.B and P1.C; and (ii) P2.A or P2.B (but not both) hold valid. If any situation other than 'YES' and 'MIXED' occurs, that will be concluded as 'NO'.
NOTE: *(i) Please refer to Section 5.6 for detailed notes on Panel unit root and stationarity tests and the two Panel cointegration tests.* *(ii) In the forthcoming sections of the chapter, individual country study as well as for pooled data analysis, the empirical estimates highlighted in 'red' indicate the case of 'NO' for the valid existence of BS effect.*

5.3 Salient Features of Model Variables and Sample Data Set

The section briefly explains the components of the preceding sections; participant variables and key features of the sample data set involved in empirical validation of the BS hypothesis. A very detailed discussion on these topics is carried in Chapter 3. However, to recap the key features I am re-stating some of the important information here.

For each country study, the competitive position of home and U.S. in international goods market is measured through three variants of rer, the regressand in my BS model. All of these three variants contain dominant shares of nontradable goods and services. These rer measures include CPI based, GDP deflator (aggregate) based and GDP Deflator (nontradables only) based rer. All three rer indices are constructed using year 2005 constant market prices.

The only regressor in my model is relative sectoral productivity gap between home (Asian countries) and U.S. ($\tilde{a}$). In order to account for the productivities of traded and nontraded sectors across countries, I proxy the variable with Average Sectoral Productivity of Labour (ASPL) in traded and nontraded sectors. ASPL is the amount of real output produced by each unit of employed labour. The real (sectoral) output is measured by real value added, measured at 2005 market prices.

The BS hypothesis is empirically tested by disaggregating the economy (both home and foreign) into four broad sectors, i.e., agriculture, manufacturing, industry, and services. For both home and U.S., agriculture and manufacturing are consistently treated as tradable sectors and industry and services together categorized as nontradables The four broad categories are constructed by aggregating seven fine categories of real economy (see Chapter 3 for details).

The sample data period ranges from 1970 to 2013. The start and end dates vary from country to country.

5.4 Model Country Study: Korea

This section begins by testing the orders of integration of the Korea/U.S. real exchange rate and the inter-country sectoral productivity differential series. I will use two unit root tests; the Augmented Dickey-fuller test and the Dickey-Fuller GLS test for this purpose.

The inclusion of a time trend in the ADF regression equation has a trade-off. While it makes allowance for trending behaviour if the variable is trend stationary, it results in a loss of power in the unit root test if the series is difference stationary. Enders (2010) suggests testing to determine if the time trend belongs in the ADF specification. If the appropriate test indicates that the time trend does not belong, he suggests estimating the ADF specification with just a constant/drift term. Unfortunately, this strategy has problems when used with the software package EViews. The null hypothesis for the "Intercept only" version of the ADF test in EViews assumes that the data generating process for the series is a random walk without drift. If the series has a drift term, the resulting critical values produced by EViews will be invalid.

Accordingly, my strategy is to undertake a visual inspection of the time plot of the series. If the series demonstrates trending behaviour, I will use the ADF unit root test specification that allows for both intercept and trend. If there is no evident trending behaviour, I will use the "intercept only" specification. When in doubt, I will include the time trend.

5.4.1 Determining the Order of Integration of rer and $\tilde{a}$:

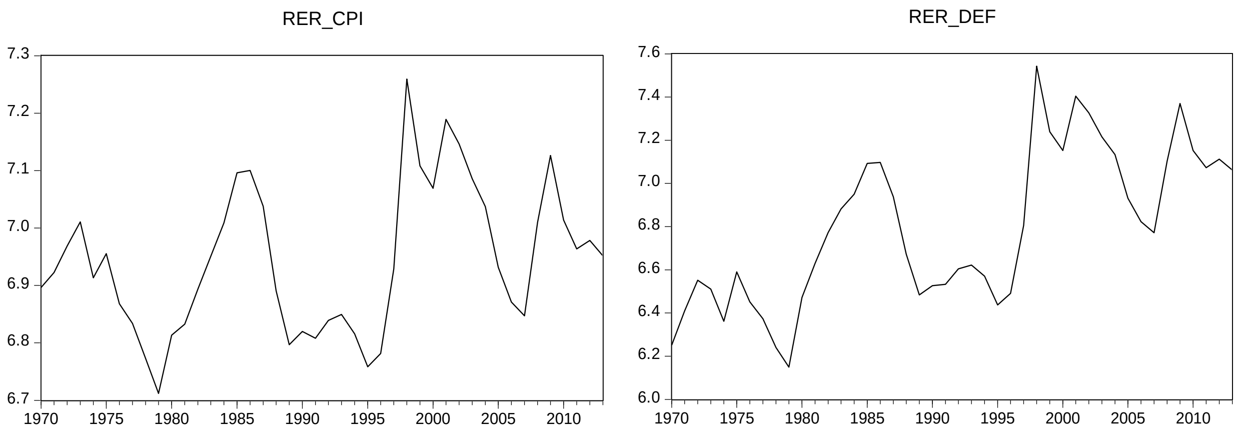
In this section, I will elaborate my procedure in detail to explain how I determined the order of integration of the two model variables. In later country studies, I will summarize the results of my testing, rather than giving full details. I note that all my results can be obtained by running the EViews programs for Korea attached as an Appendix to this chapter.

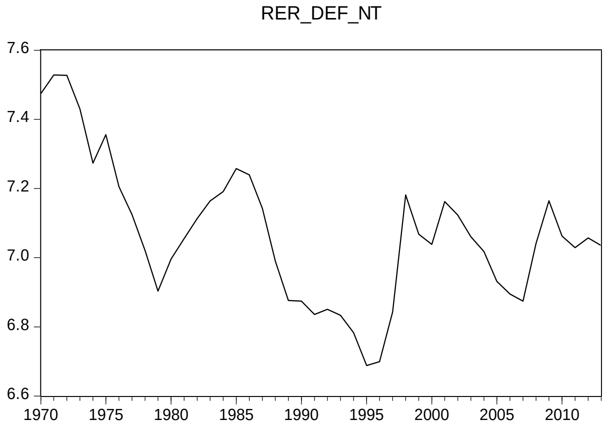
(i) Graphical Analysis of the Time-Series

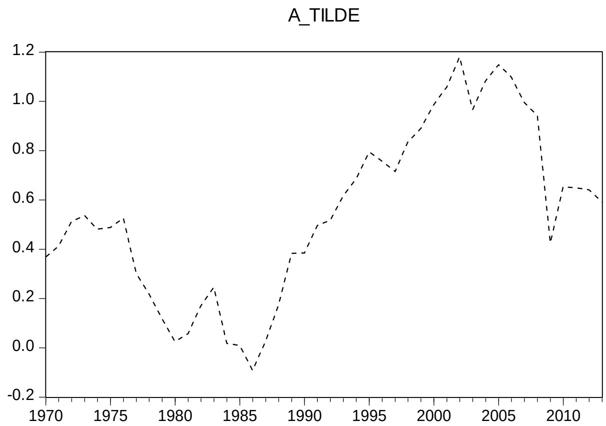
One of the most popular, informal tests for stationarity is the graphical analysis of the model time series. A visual plot of the series is usually the first (informal) step in investigating any time series before pursuing any formal tests. Also, the preliminary examination of the data is important as it allows the detection of data errors, and helps to identify structural breaks. FIGURE 5.1 plots the natural log of rer and $\tilde{\alpha}$ against time.

It is clear from FIGURE 5.1 that all the three variants of rer and $\tilde{\alpha}$ have a definite time trend and, as such, initial unit root testing should include both an intercept and a time trend in the testing specification. I also note that for most of the sample period, the rer series (except GDP deflator nontradables based version) and $\tilde{\alpha}$ are trending in similar directions, a behavior consistent with the BS hypothesis.

FIGURE 5.1: Plots of rer and $\tilde{\alpha}$ (1970-2013)







(ii) Unit Root Tests Results

The two unit root tests employed in this study are the Augmented Dickey-fuller (ADF) test and the Dickey-Fuller GLS (DF-GLS) test. The two tests commonly have the null hypothesis that a series has a unit root.

(a) <u>The Augmented Dickey-Fuller (ADF) Unit Root Test</u>

Following the ADF test specifications of equation (5.2) (inclusive of an intercept and linear time trend), I can re-write the ADF equation for rer and $\tilde{a}$ as:

(5.9-a) $$\Delta rer_t = c_{rer} + \delta_{rer} t + \emptyset_{rer} rer_{t-1} + \sum_{i=1}^{p} \beta_i \, \Delta rer_{t-i} + \varepsilon_t^{rer} \, ,$$

(5.9-b) $$\Delta \tilde{a}_t = c_{\tilde{a}} + \delta_{\tilde{a}} t + \emptyset_{\tilde{a}} \tilde{a}_{t-1} + \sum_{i=1}^{p} \gamma_i \, \Delta \tilde{a}_{t-l} + \varepsilon_t^{\tilde{a}} \, ,$$

The p lagged differenced terms of rer and $\tilde{a}$ are used to "soak up" any serial correlation in the error term that would otherwise invalidate hypothesis testing. Both error terms are assumed to be homoskedastic.

I first report the results of my ADF test, and then explain how I obtained my results. The first three columns of TABLE 5.3 below report (i) the series, (ii) the deterministic regressors, and

(iii) the number of lags included in the respective specification. Before proceeding to the unit root tests, I check whether the residuals from the respective ADF equations are serially correlated.

TABLE 5.3: ADF Unit Root Test Results for rer and $\tilde{a}$

Variables	Deterministic Regressors	Lag Length	Residual White Noise?	Sample Statistic	5% Critical Value
rer_cpi	Intercept + Trend	1	Yes	-3.32	-3.52
rer_def	Intercept + Trend	1	Yes	-3.35	-3.52
rer_def_nt	Intercept + Trend	1	Yes	-2.83	-3.52
$\tilde{a}$	Intercept + Trend	0	Yes	-1.42	-3.52

***NOTE**: Lag length is determined by automatic selection based on the SIC, subsequently adjusted to produce white noise in the residuals, if necessary.*

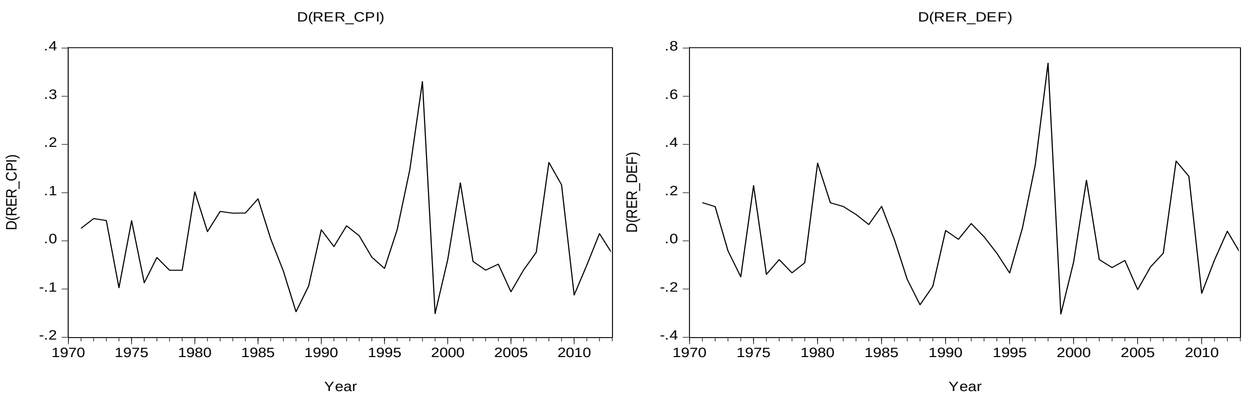
Testing for serial correlation is based on the Ljung-Box Q-statistic. The null hypothesis of the Ljung-Box Q-test is that there is no autocorrelation up to a designated number of lags, which I run from 1 to 10. For the rer_cpi series, EViews selects 1 lag (*p=1* in $\sum_{i=1}^{p} \beta_i \, \Delta rer\_cpi_{t-i}$). I find that the resulting residuals are consistent with white noise, with the Ljung-Box Q-test failing to reject the null of no serial correlation for all cumulative lags up to 10. The smallest p-value is 0.23 and occurs at cumulative lag 9. Similarly, for the rer_def series, EViews selects 1 lag. I find that the resulting residuals are consistent with white noise, with the Ljung-Box Q-test failing to reject the null of no serial correlation for all cumulative lags up to 10. The smallest p-value is 0.18 and occurs at cumulative lag 9. And finally, for the rer_def_nt series, EViews once again selects 1 lag. I find that the resulting residuals are consistent with white noise, with the Ljung-Box Q-test failing to reject the null of no serial correlation for all cumulative lags up to 10. The smallest p-value is 0.59 and occurs at cumulative lag 9.

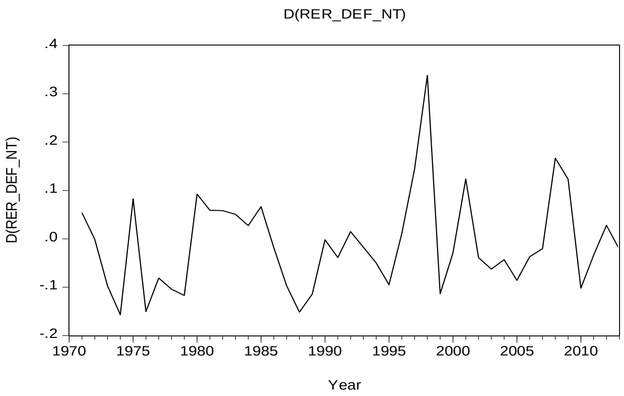
For the $\tilde{a}$ series, EViews' SIC selection algorithm also selected zero lags ($p=0$ in $\sum_{i=1}^{p} \beta_i \, \Delta \tilde{A}_{t-i}$). However, the resulting residuals showed no evidence of serial correlation. I find that the resulting residuals are consistent with white noise, with the Ljung-Box Q-test failing to reject the null of no serial correlation for all cumulative lags up to 10. The smallest p-value is 0.37 and occurs at cumulative lag 6.

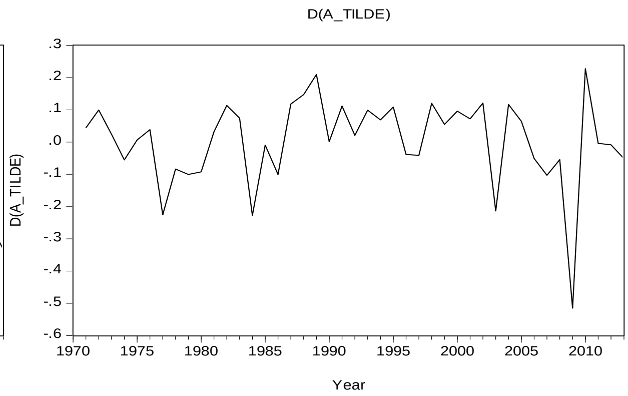
I now proceed to the respective ADF tests for unit root. The associated sample statistics at 5% critical values are reported in the last two columns of TABLE 5.3. The sample statistic for the ADF test of rer_cpi is -3.32. This is greater than the 5% critical value of -3.52. As a result, I fail to reject the null hypothesis that rer_cpi has a unit root. Similar is the situation for other two versions of real exchange rate. rer_def and rer_def_nt respectively produce sample statistics of -3.35 and -2.83 which is greater than corresponding 5% critical value of -3.52. Coming towards $\tilde{a}$, the ADF unit root test produces a sample statistic of -1.42, which is also greater than the 5% critical value of -3.51. Based on these results, I fail to reject the null that the three variants of rer $\tilde{a}$ series are unit root in levels.

Having determined that the series have unit roots, I next difference the two series and test if the differenced series are non-stationary. The time plots of the two differenced series are reported in FIGURE 5.2.

FIGURE 5.2: Plots of Δrer and $\Delta\tilde{a}$ (1970-2013)







There is no visual evidence of a decreasing or increasing trend in all three measures of Δrer and $\Delta\tilde{a}$. The series are clearly fluctuating around their natural means, with no visible trend movements. This calls for excluding trend from the deterministic regressors of my ADF test specifications. Thus, the corresponding ADF specifications for Δrer and $\Delta\tilde{a}$ are given by:

(5.10-a) $$\Delta(\Delta rer_t) = c_{rer} + \emptyset_{rer}\Delta rer_{t-1} + \sum_{i=1}^{p} \beta_i \, \Delta(\Delta rer_{t-i}) + \varepsilon_t^{\Delta rer} ,$$

(5.10-b) $$\Delta(\Delta\tilde{a}_t) = c_{\tilde{a}} + \emptyset_{\tilde{a}}\widetilde{\Delta a}_{t-1} + \sum_{i=1}^{p} \gamma_i \, \Delta(\Delta\tilde{a}_{t-l}) + \varepsilon_t^{\Delta\tilde{a}} ,$$

The results for ADF test for three versions of Δrer and $\Delta\tilde{a}$ are reported in TABLE 5.4.

TABLE 5.4: ADF Unit Root Test Results for Δrer and $\Delta\tilde{a}$

Variables	Deterministic Regressors	Lag Length	Residuals White Noise?	Sample Statistic	5% Critical Value
Δrer_cpi	Intercept	0	Yes	-5.42	-2.93
Δrer_def	Intercept	0	Yes	-5.49	-2.93
Δrer_def_nt	Intercept	0	Yes	-5.16	-2.93
$\Delta\tilde{a}$	Intercept	0	Yes	-6.51	-2.93

***NOTE**: Lag length is determined by automatic selection based on the SIC, subsequently adjusted to produce white noise in the residuals, if necessary.*

As before, I estimate the specifications with lags chosen by EViews' SIC selection algorithm, and then test the residuals for white noise. For Δrer_cpi, EViews selects zero lags ($p=0$). Applying the Ljung-Box Q-test to the resulting residuals produces the conclusion that the residuals are white noise. For cumulative lags 1 to 10, the smallest p-value is 0.34 (for cumulative lag 2). Similarly, for both Δrer_def and Δrer_def_nt, EViews once again selects zero lags. Applying the Ljung-Box Q-test to the resulting residuals reveals that the residuals are white noise. For cumulative lags 1 to 10, the smallest p-value for Δrer_def is 0.26 and for Δrer_def_nt this value is 0.58 (at cumulative lag 2 for both *rers*). Hence, the three Δrer series are pretty clean and do not show any evidence of serial correlation.

For $\Delta\tilde{a}$, EViews selects 0 lags. With 0 lags, there is no more evidence of serial correlation in the $\Delta\tilde{a}$ series. All of the p-values associated with the Ljung-Box Q-test are greater than 5%, with the smallest p-value occurring at cumulative lag 6 (p-value = 0.33). Thus, with zero lag, the Ljung-Box Q-test fails to reject the null hypothesis of no serial correlation for all cumulative lags up to 10.

I now run ADF unit root tests for Δrer and $\Delta\tilde{a}$ with 0 lags. The sample statistics for the ADF test of Δrer_cpi, Δrer_def and Δrer_def_nt are -5.42, -5.49 and -5.16 respectively. This is much smaller than the 5% critical value of -2.93. As a result, I reject the null hypothesis that any of the three variants of Δrer has a unit root. Thus Δrer is first differenced stationary, for all three variants. Similarly, the ADF unit root test for $\Delta\tilde{a}$ produces a sample statistic of -6.51, which is substantially smaller than the 5% critical value of -2.92. Based on this result, I fail to accept the null that $\Delta\tilde{a}$ has a unit root. Putting all these results together, I conclude, on the basis of the ADF tests that the level series, rer and $\tilde{a}$ are both integrated of order one, I (1).

(b) The Dickey-Fuller Generalized Least Squares (DF-GLS) Unit Root Test

The DF-GLS unit root test is described in Elliot et al. (1996). Like the ADF test, DF-GLS test also follows the null hypothesis of unit root. Following the ADF treatment above, I choose the "Intercept + Trend" and "Intercept only" specification in the subsequent testing of rer and $\tilde{a}$ and Δrer and $\Delta\tilde{a}$. TABLE 5.5 displays the DF-GLS unit root test results for the rer and $\tilde{a}$ ˜series.

TABLE 5.5: DF-GLS Unit Root Test Results for rer and $\tilde{a}$

Variables	Deterministic Regressors	Lag Length	Residuals White Noise?	Sample Statistic	5% Critical Value
rer_cpi	Intercept + Trend	1	Yes	-3.39	-3.19
rer_def	Intercept + Trend	1	Yes	-3.43	-3.19
rer_def_nt	Intercept + Trend	1	Yes	-2.51	-3.19
$\tilde{a}$	Intercept + Trend	0	Yes	-1.48	-3.19

NOTE*: Lag length is determined by automatic selection based on the SIC, subsequently adjusted to produce white noise in the residuals, if necessary.*

The associated sample statistics generated by DF-GLS test and the 5% critical values are reported in the last two columns of TABLE 5.5. The sample statistic for the DF-GLS test of rer_cpi is -3.39. This is smaller than the 5% critical value of -3.19. As a result, I may reject the null hypothesis that rer_cpi has a unit root. Thus, DF-GLS proves the rer_cpi series to be integrated of order zero, I (0). Similar to rer_cpi, the rer_def yields a sample statistic of -3.48, rejecting the null hypothesis of unit root at better than 5% significance level. Unlike the earlier two versions of real exchange rate, rer_def_nt produces sample statistics of -2.51 which is greater than corresponding 5% critical value of -3.19. Hence I may not reject the null hypothesis of unit root for GDP deflator (nontradables) based measure of rer. Coming towards $\tilde{a}$, the DF-GLS unit root test produces a sample statistic of -1.48, which is greater than the 5% critical value of -3.19. Based on these results, I also fail to reject the null of unit root that for $\tilde{a}$ series.

Next, I difference the rer_def_nt and $\tilde{a}$ series and test the status of their order of integration.

TABLE 5.6: DF-GLS Unit Root Test Results for Δrer and $\Delta\tilde{a}$

Variables	Deterministic Regressors	Lag Length	Residuals White Noise?	Sample Statistic	5% Critical Value
Δrer_def_nt	Intercept	0	Yes	-4.93	-1.94
$\Delta\tilde{a}$	Intercept	0	Yes	-6.53	-1.94

***NOTE**: Lag length is determined by automatic selection based on the SIC, subsequently adjusted to produce white noise in the residuals, if necessary.*

The test results for DF-GLS test for Δrer_def_nt and $\Delta\tilde{a}$ are given in TABLE 5.6. The two series commonly pick 0 lags through SIC automatic lag selection. The sample statistic for DF-GLS test for Δrer_def_nt is -4.93 respectively. The value is substantially smaller than the 5% critical value of -1.94. As a result, I reject the null hypothesis that Δrer_def_nt has a unit root. Thus the series is first differenced stationary. Similarly, the DF-GLS unit root test for $\Delta\tilde{a}$ produces a sample statistic of -6.53, which is smaller than the 5% critical value of -1.94. Based on this result, I reject the null that $\Delta\tilde{a}$ has a unit root. Putting all these results together, I conclude, on the basis of the DF-GLS test that the level series rer_def_nt and $\tilde{a}$ are integrated of I (1).

However, the contrasting results of ADF test and DF-GLS test leave me indecisive to conclude the true order of integration of the rer_cpi rer_def series. TABLE 5.7 summarizes the results I obtained using the two unit root tests. The last column of the table states my conclusion about the stationary status of rer and $\tilde{a}$ for Korea.

TABLE 5.7: Summary of Unit Root Test Results for rer and $\tilde{a}$

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(0)**	Inconclusive
rer_def	Yes	I(1)***	I(0)**	Inconclusive
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

5.4.2 Determination of Cointegration between rer and $\tilde{a}$

Next, I proceed towards establishing whether a long-run cointegrating relationship exists between the two series. Cointegration holds practical economic implications. While time series may be non-stationary in levels, they may move together in the long-run. A cointegrating relationship can also be considered as a long term or equilibrium phenomenon where the subject variables diverge from equilibrium in the short-run but maintain an economically valid association in the long-run.

In Section 5.2 of this chapter, I have discussed the econometric framework I will use to detect the plausible long-run BS effect for the subject Asian countries. For individual country study, two distinct cointegration approaches will serve the purpose; (i) a residual-based single equation cointegration approach, and (b) a maximum likelihood-based multivariate cointegration approach. The two tests evolve in several steps, I have elaborated these in Section 5.2.1 of the chapter. Let us start with the residual-based single equation cointegration test. Similar to my discussion on unit root testing, I will elaborate the cointegration test for Korea in a sequential manner, detailing every step involved. Once the procedure is established, the remainder of the country studies will be done similarly, but discussed with brevity. The EViews program codes for the single equation cointegration test for Korea can be found in the appendix to this chapter.

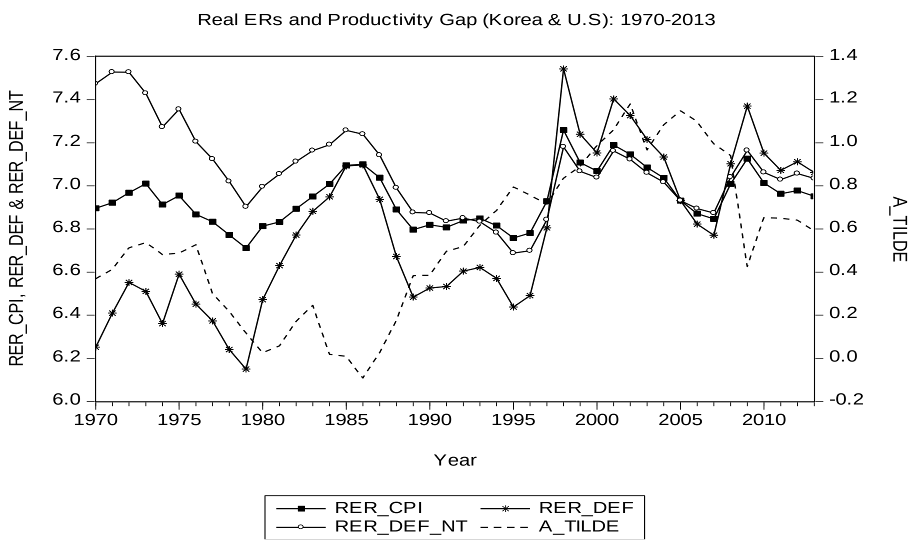
(i) Residual-Based Single Equation Cointegration Test

For the single equation cointegration test, I will use the test specification of Engle and Granger (1987), with the only regression variable being $\tilde{a}$ (between Korea and U.S.). The residuals from this regression will be tested to determine if there is a long-run association between the model's variables.

(a) <u>STEP 0: Graph the Suspected Cointegrating Series</u>

Graphing the time plots of rer and $\tilde{a}$ (between Korea and U.S.) is an informal procedure to analyse the two series for their plausible long-run association. Though the measure is informal, understanding the trend movements of two series and their adjustments over time may be helpful for obtaining a preliminary idea about their causal relationship.

FIGURE 5.3: Plots of rer and $\tilde{a}$ (1970-2013)



It is clear from FIGURE 5.3 that all three variants of rer and $\tilde{a}$ are trending in a similar direction, pointing to a (plausible) positive long-run association. As noted above, this is consistent with the BS effect. Furthermore, the successive divergence and convergence of the two type of series is consistent with cointegrating behaviour.

(b) STEP 1.A: Establishing Cointegration between rer and $\tilde{a}$

The first formal step estimates the static cointegrating relationship between rer and $\tilde{a}$. All dynamics are ignored and the cointegrating regression is estimated by the OLS.

(5.3)' $\epsilon_t = rer_t - \hat{\alpha} - \hat{\beta}\, \tilde{a}_t$,

where the OLS residuals (ϵ_t) are a measure of disequilibrium. In order to establish cointegration between rer and $\tilde{a}$, the necessary condition is that the estimated residuals from Eq. (5.3)' should be stationary, (i.e., $\epsilon_t \sim$ I (0)). Since rer and $\tilde{a}$ are non-stationary (which causes the famous 'spurious regression problem'), one should place little faith in the standard error estimates (and thus t-statistics) in the cointegrating regression. So, the true order of integration will be judged by comparing the test statistics with Mackinnon (1996) critical values.

(5.4)' $\Delta\epsilon_t = \delta + \varphi\epsilon_{t-1} + \sum_{j=1}^{p} \omega_j\, \Delta\epsilon_{t-j} + u_t$.

Under the condition of mean reverting residuals, I (0), the estimated statistics for φ will be compared against the MacKinnon (1996) critical values at the 10% significance level. The test is run by EViews using Engle-Granger test specifications. TABLE 5.8 reports the test results for STEP 1.

The EG test results reveal that only the cointegration regression of rer_cpi and $\tilde{a}$ is consistent with a long-run association. The model rejects the null hypothesis of no cointegration at better than the 5 percent significance level. This implies that the model residuals display short-lived memory and are mean reverting. However, the other two models could not reject the null of no cointegration, even at a 10 percent significance level. Thus, I proceed towards the next set of estimations for establishing BS effect only for the series rer_cpi and $\tilde{a}$.

TABLE 5.8: Testing the Cointegration between rer and $\tilde{a}$ Using Engle-Granger Cointegration Test

$rer_t = \alpha + \beta\, \tilde{a}_t + \epsilon_t$

Null Hypothesis: rer and $\tilde{a}$ are not cointegrated (ϵ_t does not hold a short memory)

Dependent Variable	Deterministic Regressors	Lags	Tau-Statistic	*p*-value	Are Residuals (ϵ_t) I(0)?
rer_cpi	None	1	-3.55	0.04	Yes
rer_def	None	0	-2.91	0.15	No
rer_def_nt	None	0	-2.30	0.38	No

NOTE: *Lag length is determined through SIC, subsequently adjusted to produce white noise in the residuals, if necessary.*

(c) *STEP 1-B: Estimating the Error Correction Model (ECM)*

The second part of STEP-1 consists of estimating a short-run model with an error-correction mechanism (ECM), using a Newey-West (NW) HAC OLS estimator. According to the Granger Representation Theorem, if rer and $\tilde{a}$ are cointegrated, then there will exist an ECM relating these variables and vice versa. After estimating the three variants of the BS model through the EG residual based tests, I now check the proposed long-run relationship between rer and $\tilde{a}$ through the ECM. For this purpose, the long-run regression (5.3)' may be used in the following short-run model, with the remaining parameters being consistently estimated by the NW HAC OLS estimator.

$$(5.5)' \quad \Delta rer_t = \gamma + \rho\left(rer_{t-1} - \hat{\alpha} - \hat{\beta}\, \tilde{a}_{t-1}\right) + \sum_{i=1}^{k} \mu_i\, \Delta rer_{t-k} + \sum_{i=1}^{k} \lambda_i\, \Delta\, \tilde{a}_{t-k} + \nu_t\ ,$$

I am using the augmented version of the ECM, where difference-lagged terms of rer and $\tilde{a}$ are allowed to contribute to the short-run dynamics of the model. Note that the estimated coefficient ρ in the short-run Eq. (5.5)' should have a negative sign and be statistically significant. Note also that, to avoid an explosive process, the coefficient should ideally take a value between 0 and -1.

In my analysis, STEP 1.B of the single equation cointegration test involves several parts. The incorporation of difference-lagged terms of rer and $\tilde{a}$ in ECM requires me to specify an appropriate number of lags for the two series. Since rer and $\tilde{a}$ series are taken on an annual basis, the lag order selection is done from a maximum of 4 lags in order to ensure good levels of adjustment in the model and for the attainment of well-behaved residuals.

Selecting the Number of Lags for rer

The appropriate number of difference-lagged terms of rer_cpi, rer_def and rer_def_nt in their respective error correction models will be decided through a VAR model, using four different information criteria. From TABLE 5.9, we may see that the four selection criteria are unanimously selecting one lag for three different VAR models of rer and $\tilde{a}$. However, in the event of conflicting suggestions made by different selection criteria, I follow the lag selection proposed by SIC.

The lag length test above indicates that there should be one lag of rer_cpi, rer_def and rer_def_nt while estimating their respective EC regressions. But before running the ECM regressions, I tested the VAR residuals from the three models for serial correlation through the LM test. The residuals turned out to be absolutely white only for the case of the rer_cpi based model. So, I am satisfied with the selection of one lag for the rer_cpi based ECM, as indicated by the four information criteria. However for the cases of rer_def and rer_def_nt, the VAR residuals did not turn out to be white noise at one lag. As a result, I increased the number of lags from one through four and at four lags I obtained white residuals.

TABLE 5.9: Information Criteria Values for Different Lag Lengths for the Multivariate VAR for rer and $\tilde{a}$

Information Criteria	Lags for rer_cpi				
	0	**1**	**2**	**3**	**4**
FPE	0.00	0.00*	0.00	0.00	0.00
AIC	-0.39	-3.02*	-2.97	-2.80	-2.72
SC	-0.31	-2.77*	-2.55	-2.21	-1.96
HQ	-0.36	-2.93*	-2.82	-2.59	-2.45
Information Criteria	**Lags for rer_def**				
	0	**1**	**2**	**3**	**4**
FPE	0.01	0.00*	0.00	0.00	0.00
AIC	1.45	-1.46*	-1.37	-1.21	-1.06
SC	1.54	-1.21*	-0.94	-0.62	-0.30
HQ	1.48	-1.37*	-1.21	-1.00	-0.78
Information Criteria	**Lags for rer_def_nt**				
	0	**1**	**2**	**3**	**4**
FPE	0.00	0.00*	0.00	0.00	0.00
AIC	-0.03	-2.93*	-2.85	-2.68	-2.57
SC	0.00	-2.67*	-2.43	-2.09	-1.81
HQ	-0.00	-2.83*	-2.69	-2.47	-2.30

** Indicates lag order selected by the criterion*
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion

<u>Selecting the Number of Lags for $\tilde{a}$</u>

For selecting the appropriate number of lags of $\tilde{a}$ in my EC regressions, I run a preliminary error correction model, with 0, 1 and 2 lags of $\tilde{a}$ respectively in the search of obtaining the most efficient specification of ECM, generating white residuals. I use the Schwarz Information Criterion (SIC) to determine the appropriate number of lags.

TABLE 5.10 reports the SIC values for the preliminary EC regressions, including 0, 1 and 2 lags of $\tilde{\alpha}$ respectively.

TABLE 5.10: Selecting the Appropriate Number of Differenced-Lags of $\tilde{\alpha}$ for ECM

Information Criteria	Number of Differenced-Lags of $\tilde{\alpha}$ in ECM		
	0	**1**	**2**
$\boldsymbol{SIC_{rer\_cpi}}$	-1.96*	-1.87	-1.76
$\boldsymbol{SIC_{rer\_def}}$	-0.15*	-0.077	0.01
$\boldsymbol{SIC_{rer\_def\_nt}}$	-1.60*	-1.53	-1.44

NOTE

*(i) * Indicates lag order selected by the SIC criterion.*
(ii) One lag of rer_cpi and four lags of rer_def and rer_def_nt are included in their respective preliminary EC regressions.

For all three models, SIC is reporting its least value where no difference-lag terms of $\tilde{\alpha}$ are included. Accordingly, I put zero lags of $\Delta\tilde{\alpha}$ in my EC estimations and tested the models' residuals for serial correlation. The residuals were not white. I raised the number of lags (of both rer and $\tilde{\alpha}$) but still could not produce residuals free from serial correlation. So, I decided to stick to my original EC specification; i.e., using one difference-lagged term for rer_cpi, four difference-lagged terms for rer_def and rer_def_nt and no lagged values of $\Delta\tilde{\alpha}$.

TABLE 5.11 reports the ECM results, for the three test regressions of rer and $\tilde{\alpha}$ (Equation (5.5)'). The ECM model produces statistically significant error correction coefficients of -0.38, -0.99 and -0.63 for the rer_cpi, rer_def and rer_def_nt based models, respectively. Interpreting the rer_cpi based model results, the series adjusts so that 38 percent of the deviations from long-run equilibrium is "corrected" each period. The EC coefficient values of the rer_def and rer_def_nt based EC models can be interpreted analogously. Thus, the respective ECMs display significant error correction processes, a finding endorsing my rer_cpi model results (only) yielded through EG test. However, with respect to the rer_def and rer_def_nt based models, the two cointegration tests generate contrasting results. However, as my criteria only require that one of the cointegration tests indicate that the series are cointegrated (at least one of S1.A and S1.B), I proceed to estimate long-run relationships for all three measures of real exchange rates.

TABLE 5.11: Error Correction Model (ECM) Results for rer and $\tilde{a}$

$$\Delta rer_t = \gamma + \rho(rer_{t-1} - \alpha - \beta\ \tilde{a}_{t-1}) + \sum_{i=1}^{k} \mu_i\, \Delta rer_{t-k} + \nu_t\ ,$$

Dependent Variable	Lags of rer	Error Correction (EC) Coefficient	Does Significant Speed of Adjustment Hold?
rer_cpi	1	-0.38 [-3.38]	Yes
rer_def	4	-0.99 [-3.05]	Yes
rer_def_nt	4	-0.63 [-2.94]	Yes

NOTE
(i) t-values are given in squared-brackets.

(d) *<u>STEP 2: Estimating the Long-Run Balassa-Samuelson Coefficient</u>*

Having cointegration established between rer and $\tilde{a}$ through the EG test and/or the ECM, the second and last step of investigating the BS hypothesis is to estimate the long-run BS coefficient of the model. For this purpose, I employ Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS) single equation cointegration estimators. The two estimators yield varying results for the three versions of the BS model. For the case of the rer_cpi based model, the DOLS results indicate that the long-run BS coefficient (β) holds a positive value, though with weak statistical significance. This implies that trend movements in relative sectoral productivity differences between Korea and U.S. are significantly driving CPI based rer appreciation in Korea. In contrast, FMOLS test yields a positive but statistically insignificant BS coefficient. These contradictory results on the part of FMOLS and DOLS estimators makes me conclude that the empirical evidence for the BS effect for Korea is “Mixed”. On the other hand, for rer_def and rer_def_nt based models, the two estimators yield common results. For the case of rer_def based model, both DOLS and FMOLS tests indicate a valid BS effect for the country. The two estimators produce positive and statistically significant (at better than 5 percent significance level) long-run BS coefficients. However, the two estimators generate negative and statistically insignificant BS coefficients when the hypothesis is tested using rer_def_nt and $\tilde{a}$. This makes me conclude “NO” for the BS hypothesis when the variable rer_def_nt is used to measure the real exchange rate for Korea and the U.S.

TABLE 5.12: Estimating the Long-Run BS Coefficient through FMOLS and DOLS Cointegration Regression Estimators

Dependent Variable	**Long-Run BS Coefficient (β)**		**Does BS Effect Hold?**
	FMOLS	**DOLS**	
rer_cpi	0.13 [1.58]	0.16 [1.65]	Mixed
rer_def	0.51 [2.21]	0.57 [2.35]	Yes
rer_def_nt	-0.14 [-0.99]	-0.11 [-0.80]	No

NOTE
(i) t-values are given in squared-brackets.

TABLE 5.13 summarizes the test results for all three steps (STEP 0 to 2) of the single equation cointegration test. The last column of the table states my conclusion on the valid/invalid existence of the BS effect for Korea against U.S.

TABLE 5.13: Summary of Single Equation Cointegration Test Results for BS Hypothesis

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	Yes	1	-0.38 [-3.38]	0.13 [1.58]	0.16 [1.65]	Mixed
rer_def	No	4	-0.99 [-3.05]	0.51 [2.21]	0.57 [2.35]	Yes
rer_def_nt	No	4	-0.63 [-2.94]	-0.14 [-0.99]	-0.11 [-0.80]	No

NOTE
(i) t-values are given in squared-brackets.

(ii) Maximum Likelihood Based Multivariate Cointegration Approach

Similar to the single equation cointegration method, there are a few distinct steps involved to establish whether a long-run association of rer and $\tilde{a}$ exists when using the multivariate cointegration approach. Below, I will discuss each of these steps individually.

(a) Determining the Lag Length

Before the cointegration test can be carried out, it is necessary to specify the number of lags that will be used in the associated Johansen ML cointegration as well as the VEC model. For the single equation test, I already found the appropriate number of lags for the rer_cpi, rer_def and rer_def_nt based cointegration models using VAR. The readers may refer to TABLE 5.9 and the associated discussion for details. As I already have discussed, in my efforts to eliminate serial correlation from my estimations, I have to have one lag of rer_cpi and four lags of rer_def and rer_def_nt in their respective regressions. This is indicative of the fact that I will include one and four lags in my rer_cpi, rer_def and rer_def_nt based multivariate cointegration models respectively. Thus, subsquently, I will include 0 (= 1-1) lags in rer_cpi based VEC model and 3 (= 4 - 1) lags in my rer_def and rer_def_nt based VEC regressions.

(b) <u>*Deciding the Specification of Deterministic Regressors*</u>

After determining the appropriate lag length for the multivariate cointegration test, I test for a cointegrating relationship between rer and $\tilde{a}$. As I already have mentioned in the earlier sections of this chapter, I employ VAR-based cointegration tests using the methodology developed by Johansen (1988, 1991) and Johansen and Juselius (1990).

The specification of deterministic regressors in the Johansen test is very important. EViews allows the following 5 specifications of deterministic regressors:

Case 1: Assumes no deterministic trend in the data and no intercept or trend in the VAR and in the cointegrating equation (CE)

Case 2: Assumes no deterministic trend in the data, but an intercept in the CE and no intercept in VAR

Case 3: Assumes a linear deterministic trend in the data and an intercept in CE and test VAR

Case 4: Allows for a linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Case 5: Allows for a quadratic deterministic trend in data, intercept and trend in CE and linear trend in VAR.

I choose to employ specifications 3 and 4 of the test as these allow a reasonable degree of generality in incorporating trending behaviour in the data. Thus, the existence/inexistence of cointegration between rer and $\tilde{a}$ will be decided on the test results of Case 3 and Case 4.

(c) *STEP 1: Testing for Cointegration through Maximum Likelihood Based Rank Test*

EViews uses two tests, the Trace and Maximum Eigenvalue tests to determine whether the series are cointegrated and, if they are, the number of cointegrating equations that exist.[15] Given the two series, there can either be 0, 1, or 2 cointegrating equations. A finding of 0 cointegrating vectors indicates that the series are not cointegrated. A finding of 2 cointegrating vectors indicates that the model variables are not unit root in levels. A finding of 1 cointegrating vector means that the two series are non-stationary and cointegrated.

EViews selects the number of cointegrating equations conditional on the specification of deterministic regressors included in the various model specifications of Johansen test (see Cases 1 through 5 above). In TABLE 5.14, I report the results of this analysis for Case 3 and 4, since my decision on existence of cointegration is driven by these two specifications only.

TABLE 5.14: Multivariate Cointegration Test Results

Johansen Cointegration Rank Test Results			
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

On the whole, there is no statistical evidence in support of valid cointegration between rer and $\tilde{a}$. This is true for all three types of models, under both specifications of the Johansen test. I

[15] See Johansen and Juselius (1992) for details regarding these two tests.

obtain a rank zero for the Trace statistic as well as the Maximum Eigenvalue test, under both specifications of the Johansen ML test. These findings are not very compatible with my single equation test results for the rer_cpi and rer_def based models (only). Nevertheless, for rer_def_nt based model, the two tests yield consistent findings. The multivariate cointegration model results do not allow me to proceed further with VEC estimation. Thus, after examining all three versions of the BS model (rer_cpi, rer_def and rer_def_nt based cases) using the multivariate cointegration approach, I conclude that cointegration does not exist, making the BS effect unsupported by the statistical evidence from Korea.

TABLE 5.15: Cointegration Tests Results for Korea (1970-2013)[16]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(0)**	Inconclusive
rer_def	Yes	I(1)***	I(0)**	Inconclusive
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[17]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	Yes	1	-0.38 [-3.38]	0.13 [1.58]	0.16 [1.65]	Mixed
rer_def	No	4	-0.99 [-3.05]	0.51 [2.21]	0.57 [2.35]	Yes
rer_def_nt	No	4	-0.63 [-2.94]	-0.14 [-0.99]	-0.11 [-0.80]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

[16] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[17] t-values are given in squared-brackets.

5.5 Individual Country Studies

This section begins country-by-country reporting of results acquired through testing the BS hypothesis. The previous section on Korea provided a detailed report of the many steps involved in testing the BS hypothesis. The remainder of this chapter gives a much abbreviated report for each country in the interests of brevity, starting with Indonesia.

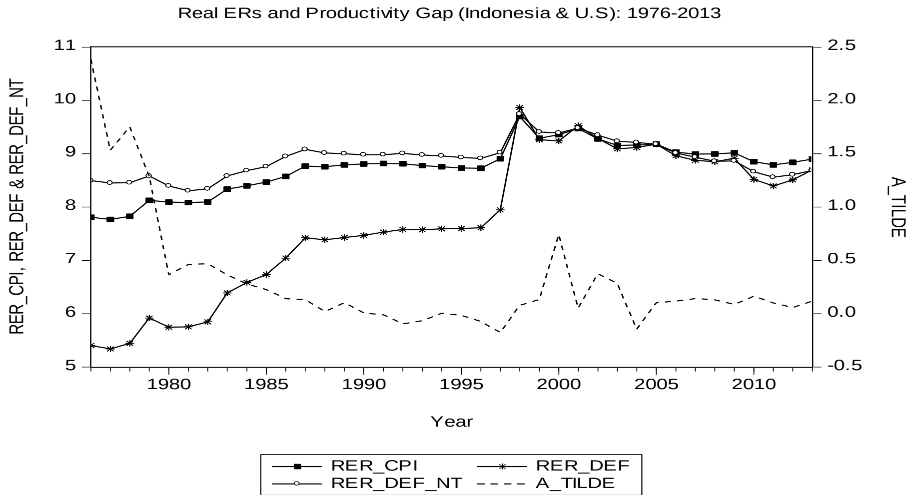
5.5.1 Indonesia

I begin by displaying the time plots of the three alternative measures of rer and $\tilde{\alpha}$ of Indonesia against U.S. in FIGURE 5.4. The visual evidence consistent with the BS hypothesis is (a) the two GDP Deflator based rer and $\tilde{\alpha}$ series are trending in a similar direction against the U.S., and (b) rer is adjusting at a modest speed to close the gap resulting from $\tilde{\alpha}$ shocks that drive the rer series away from its long-run equilibrium.

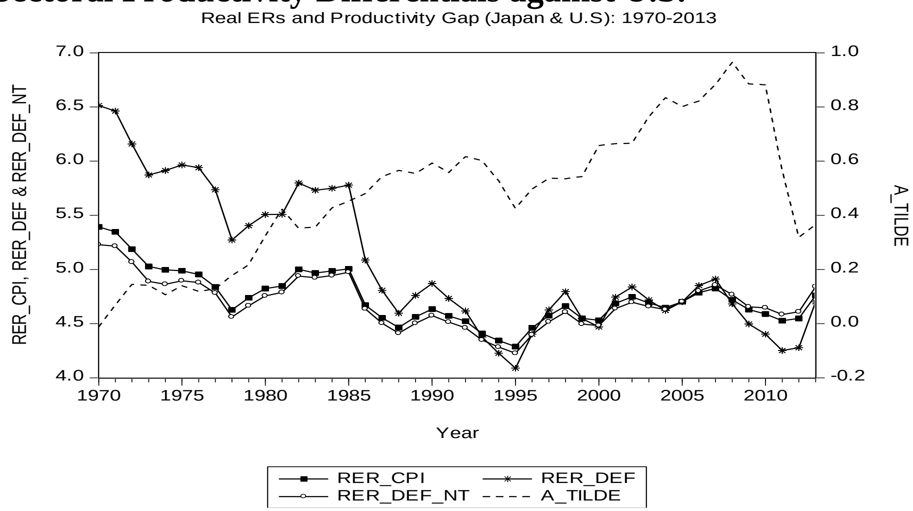
TABLE 5.16 reports empirical results for the unit root and cointegration test procedures for Indonesia. The order of integration of the three rer measures and $\tilde{\alpha}$ series are determined through the two unit root tests (ADF and DF-GLS). For all three rer series, the two tests commonly propose rer to be a unit root process in levels.

However, $\tilde{\alpha}$ series produces conflicting results. The series turn out to be level stationary, according to ADF test findings. But the DF-GLS test results suggest that the order of integration is greater than 1. This leaves me indecisive about the actual order of integration of $\tilde{\alpha}$.

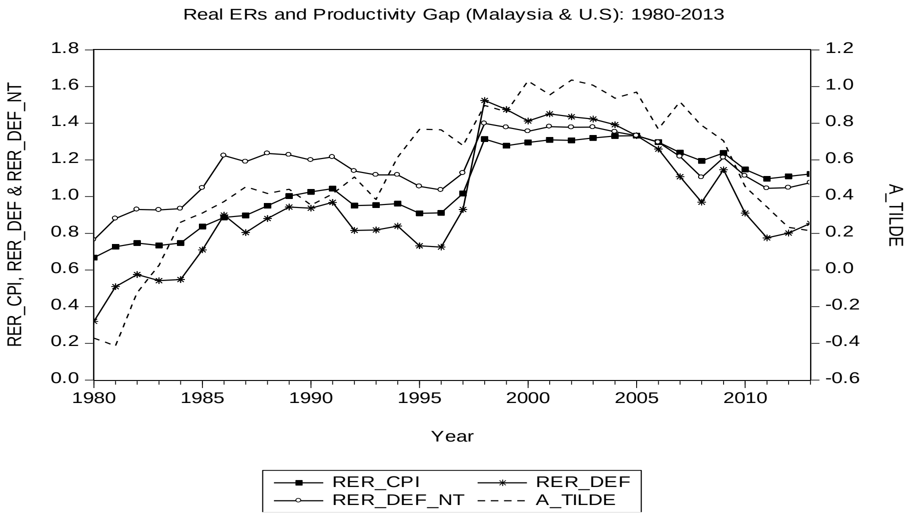
FIGURE 5.4 Plots for Real Exchange Rates and Sectoral Productivity Differentials against U.S.



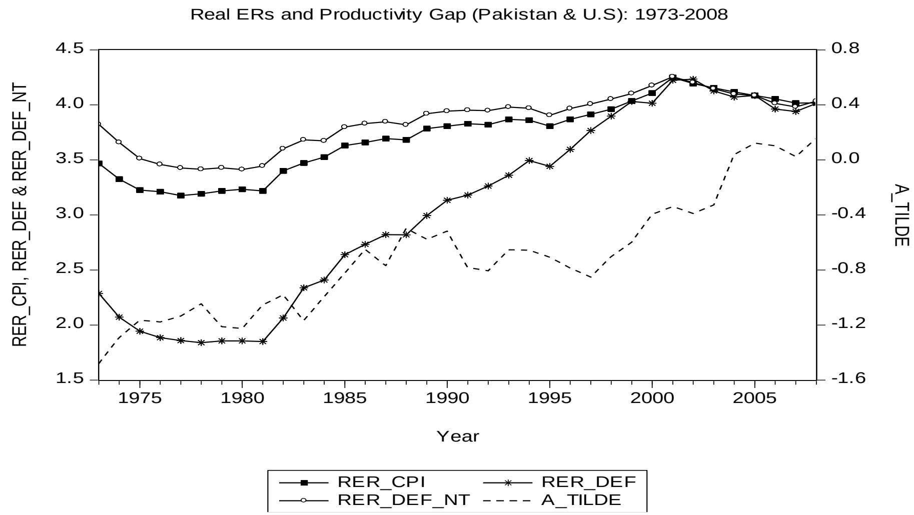
Indonesia



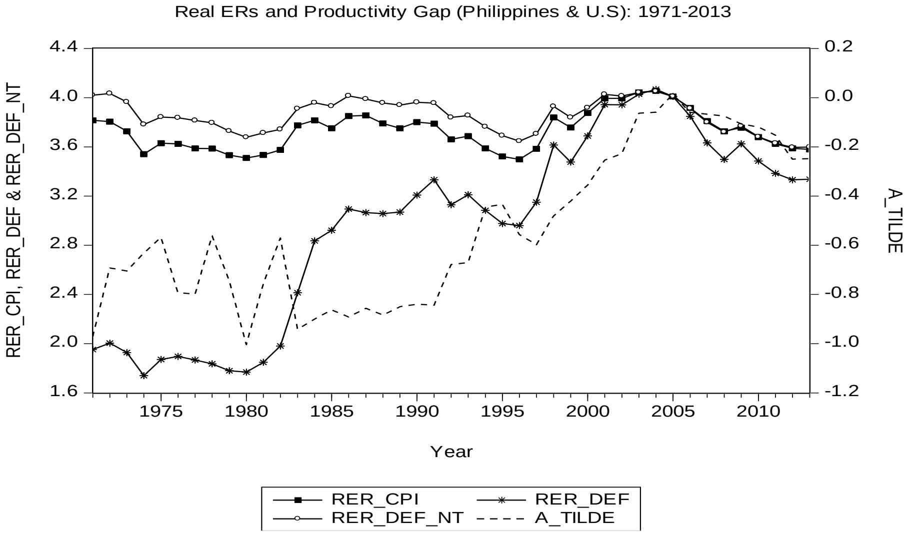
Japan



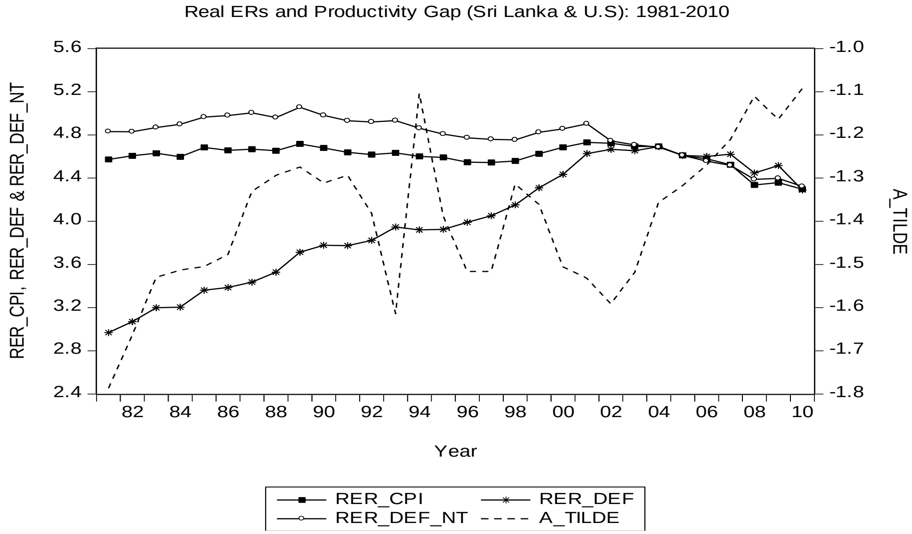
Malaysia



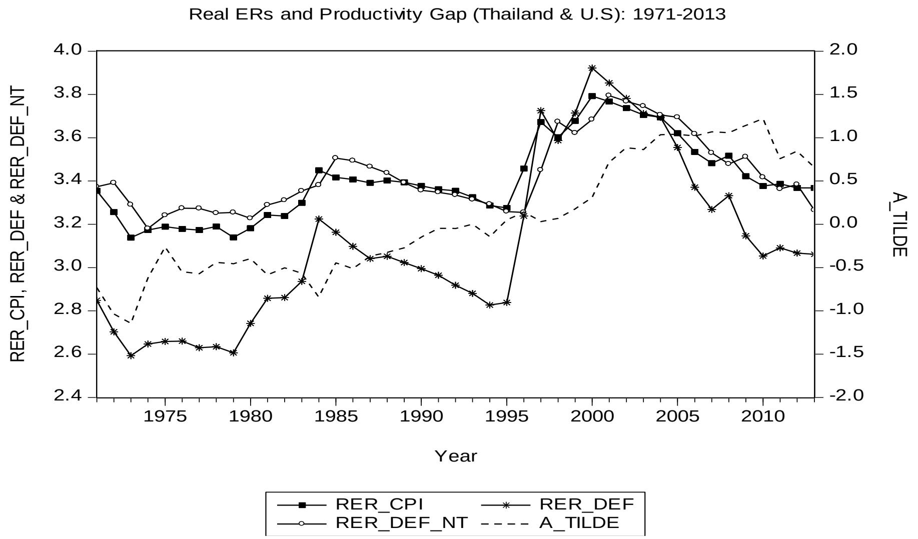
Pakistan



Philippines



Sri Lanka



Thailand

TABLE 5.16: Cointegration Tests Results for Indonesia (1976-2013)[18]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[19]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	3	-0.20	-0.50	0.36	No
			[-3.51]	[-2.29]	[0.68]	
rer_def	No	3	-0.40	-1.21	1.06	No
			[-3.36]	[-1.72]	[0.57]	
rer_def_nt	No	3	-0.20	-0.23	0.28	No
			[-3.17]	[-1.18]	[0.58]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	2	2	No
rer_def	2	2	No
rer_def_nt	2	2	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	1	1	See below
rer_def	1	1	See below
rer_def_nt	1	1	See below

Vector Error Correction Model

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
rer_cpi	2	Yes	-0.02	0.07	5.34	No
			[-1.33]	[7.39]	[5.72]	
rer_def	2	Yes	-0.00	0.02	22.15	No
			[-0.56]	[7.62]	[5.72]	
rer_def_nt	2	Yes	-0.00	0.02	21.34	No
			[-0.41]	[7.20]	[5.15]	

[18] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[19] t-values are given in squared-brackets.

From the visual inspection of the model time series, I was expecting a valid long-run association between rer and $\tilde{a}$. However, the EG single equation cointegration test results indicate otherwise. The test results reveal an absence of long-run co-movement between the model variables for all three versions of the model. The null hypothesis of no cointegration between model variables is not rejected with desired statistical precision, when tested against MacKinnon (1996) critical values using the EG cointegration test specifications. As a result, I conclude that the residuals are not mean-reverting.

In contrast, the ECM results suggest a valid cointegrating relationship. For all three model variants, the test yields a negative and statistically significant EC coefficient, suggesting significant adjustments on the part of rer to ensure its convergence to long-run equilibrium. These findings provide me with a reason to proceed further with estimating the cointegration regression (FMOLS and DOLS) estimations. However, the two cointegration regression estimators invalidate the BS hypothesis. For all three model variants, the two estimators yield either statistically insignificant or/and negative long-run BS coefficients. Thus, the single equation cointegration test does not support the BS hypothesis for Indonesia.

Compared to the single equation test results, the multivariate cointegration method yields similar findings. There is no evidence for a cointegrating relationship using the Case 3 specification of the VECM. However, the Case 4 specification suggests the presence of a cointegrating relationship. Both test statistics conclude a rank of one for Case 4. This allows me to conduct VEC estimations (see the last panel of the table). Unfortunately, the results obtained through estimation of the VEC model are rather disappointing. Two out of four model conditions, mandatory for establishing a valid BS effect are clearly violated; (a) the first EC coefficient (EC_{rer}) is statistically insignificant, even up to 10 percent significance level, implying no significant adjustments on part of rer for obtaining long-run equilibrium, and (b) the second error correction coefficient ($EC_{\tilde{a}}$) is highly statistically significant, violating the pre-condition of weak exogeneity. Thus, the multivariate cointegration approach finds no evidence for the BS effect for Indonesia, which is consistent with my findings using the single equation approach.

5.5.2 Japan

In FIGURE 5.4, the visual inspection of Japan's rer and $\tilde{a}$ differential against the U.S. does not display a very supportive environment for establishing a long-run association. This is because (a) the rer and $\tilde{a}$ series are not trending in a similar direction and, (b) the two series are rarely intersecting each other throughout the sample period.

My suspicion gets mixed empirical support on conducting the single equation cointegration test. For all the three models of rer, the single equation test (under EG specifications) rejects the null hypothesis of no cointegration with better than 10 percent statistical precision. The tau-values are tested against MacKinnon (1996) critical values, suggesting valid long-run causality between inter-country $\tilde{a}$ and rer for Japan against the U.S. Next, I move towards estimating the error correction model to know the speed of adjustment imparted by the rer series for correcting its short-term errors. The EC coefficient values range from -0.22 to -0.25; i.e., 22 to 25 percent of the each-period's fluctuations in rer are self-adjusted, thus causing the rer to return to its long-run equilibrium. But as expected, the long-run BS effect holds a negative sign, as evident from the FMOLS and DOLS estimators. Thus, Japan's inter-country productivity gap with U.S. is triggering depreciation in rer, instead of appreciation, a disagreeable behavior in the context of the BS hypothesis.

TABLE 5.17: Cointegration Tests Results for Japan (1970-2013)[20]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[21]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	Yes	2	-0.25 [-3.16]	-0.52 [-2.67]	-0.40 [-1.92]	No
rer_def	Yes	2	-0.22 [-3.33]	-1.94 [-4.92]	-1.87 [-4.27]	No
rer_def_nt	Yes	2	-0.22 [-2.65]	-0.39 [-1.84]	-0.26 [-1.13]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

[20] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[21] t-values are given in squared-brackets.

Next, I investigate the model using the Johansen ML Cointegration test procedure. For all three variants of rer, both Trace and Maximum Eigenvalue statistics support the inexistence of a valid cointegrating relationship. The two test statistics universally produce a rank of zero, for both specifications of the Johansen cointegration test. Thus, in accordance with the single equation findings, I find no support for the BS hypothesis for Japan.

5.5.3 Malaysia

From the time plots of rer and $\tilde{a}$ of Malaysia against the U.S., the plausible existence of cointegration between the two series is evident for all the three versions of rer. Throughout the sample data period, the $\tilde{a}$ series is closely co-moving with rer. Also, the two series are intersecting time and again.

The cointegration test results are provided in TABLE 5.18. Parallel to my expectations, the single equation cointegration test is suggesting the possibility of long-run co-movement between all three variants of rer and $\tilde{a}$, as evident from cointegration test run under EG residual-based test or/and ECM. The tau statistics of the EG test, when compared against MacKinnon's (1996) critical values, rejects the null hypothesis of no cointegration at better than five percent statistical significance, but only for the two GDP deflator based versions of the model. When tested for error correction model, I obtain error correction coefficients of value -0.49, -0.46 and -0.48 for the CPI, GDP deflator and GDP deflator (nontradables) based cases of the model. Finally, both the FMOLS and DOLS cointegration regression estimators produce positive and statistically highly significant long-run BS coefficients against all three variants of the BS model. Thus, according to the single equation cointegration methods, there is evidence of a BS effect for Malaysia.

TABLE 5.18: Cointegration Tests Results for Malaysia (1980-2013)[22]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[23]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of *rer*	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	4	-0.49[24] [-3.44]	0.44 [6.79]	0.54 [6.59]	Yes
rer_def	Yes	1	-0.46 [-3.19]	0.74 [6.59]	0.92 [5.17]	Yes
rer_def_nt	Yes	2	-0.48 [-2.99]	0.36 [5.73]	0.38 [3.44]	Yes

Multivariate Cointegration Approach

Version of *rer*	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	0	0	No
rer_def	2	2	No
rer_def_nt	0	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	0	0	No
rer_def	2	0	No
rer_def_nt	0	0	No

[22] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[23] t-values are given in squared-brackets.
[24] The test regression also contains first lagged-difference of $\tilde{a}$.

Contrary to the single equation cointegration model results, the BS hypothesis is not supported by the multivariate model. The Trace and Maximum Eigenvalue statistics of Johansen ML test together indicate a rank of zero and/or two, reflecting no long-run association between $\tilde{a}$ and three variants of rer.

5.5.4 Pakistan

The time movements of Pakistan's rer and $\tilde{a}$ against the U.S. (FIGURE 5.4) appear favourable for establishing a valid BS effect for the country. The $\tilde{a}$ series is closely co-moving with the two GDP deflator based rer variables. However, the rare intersection of the two series points against the existence of a long-run relationship.

Focusing on the single equation cointegration test results, the evidence of a long-run association between rer and $\tilde{a}$ is mixed. For two out of three variants of the model, the EG and ECM tests fail to establish cointegration between the model variables with acceptable statistical significance. The only evidence comes from the case of the GDP deflator (nontradables) based model where the ECM test (only) yields a negative and statistically significant EC coefficient (though with relatively weak statistical significance). The FMOLS and DOLS estimator results suggest valid existence of a long-run BS effect for the country, as the long-run BS coefficient generated through two estimators is positive with high statistical significance.

The results obtained through the multivariate model are similar to earlier ones. The Trace and Maximum Eigenvalue statistics under both cases of the Johansen ML cointegration test (Case 3 and Case 4) produce a rank of zero for all three estimated models. This demonstrates inexistence of a long-run association between $\tilde{a}$ and rer movements. On the whole, I conclude that there is a lack of empirical support in favour of a valid BS effect for Pakistan.

TABLE 5.19: Cointegration Tests Results for Pakistan (1973-2008)[25]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)**	I(1)**	I(1)
rer_def_nt	Yes	Greater than I(1)	I(1)**	Inconclusive
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[26]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient[27]	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	1	-0.05 [-0.98]	-	-	No
rer_def	No	1	-0.08 [-0.84]	-	-	No
rer_def_nt	No	3	-0.11 [-1.85]	0.51 [4.27]	0.69 [5.05]	Yes

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

[25] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[26] t-values are given in squared-brackets.
[27] The three test regressions also contains first and second lagged-difference of $\tilde{a}$.

5.5.5 Philippines

The time plots of Philippines' rer and $\tilde{a}$ series are suggestive that a BS effect exists. In FIGURE 5.4, one may see that over time the $\tilde{a}$ series has trended upwards. The series shows co-movements with the three versions of rer for parts of the sample period. From the year 1971-2005, the series closely co-move with the GDP Deflator based rer. From 2005-13 it moves along with the CPI and GDP deflator (nontradables) based rer. These patterns can be taken as informal support for cointegration between the two series.

My analysis from the preliminary visual inspection gets confirmation on conducting single equation cointegration tests. The cointegration model of CPI based rer and $\tilde{a}$ yields a valid BS effect for Philippines. Through the EG test, the model residuals turn out to be mean reverting and the tau statistic rejects the null of no cointegration at a satisfactory significance level (a little above 5 percent). A highly significant EC coefficient value of -0.39 confirms that rer movements adjust to return the series back to its long-run equilibrium. And finally, the long-run BS coefficient is estimated to be positive and significant using both the FMOLS and DOLS cointegration regression estimators. This provides support for the hypothesis that appreciation of the Philippine exchange rate is significantly determined by movements in the country's sectoral productivity gap against the U.S.

The GDP deflator based version of the model also reveals a valid long-run association between rer and $\tilde{a}$, as verified through the ECM results (only). The subsequent long-run BS coefficient (obtained through FMOLS and DOLS estimators) confirm the ECM findings by showing a (statistically) highly significant and positive BS effect holds for the country. However, the *rer_def_nt* results do not support the BS hypothesis.

TABLE 5.20: Cointegration Tests Results for Philippines (1971-2013)[28]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(0)***	I(1)***	Inconclusive
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[29]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	Yes	4	-0.39 [-3.14]	0.20 [1.65]	0.23 [1.65]	Yes
rer_def	No	1	-0.36 [-2.22]	1.84 [3.59]	1.70 [3.60]	Yes
rer_def_nt	No	3	-0.27 [-2.63]	-0.06 [-0.62]	-0.03 [-0.27]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

[28] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[29] t-values are given in squared-brackets.

A different story emerges when I use the multivariate approach. I consistently obtain a rank of zero for all three cases of rer and $\tilde{a}$, as is evident from both Trace and Maximum Eigenvalue statistics for the Case 3 and Case 4 specifications of the Johansen ML model. I conclude that the evidence for the BS hypothesis for the Philippines is mixed, with different results emerging from the single and multiple equation frameworks.

5.5.6 Sri Lanka

FIGURE 5.4 plots the model series for providing visual evidence of cointegration for Sri Lanka. The rer and $\tilde{a}$ series are not co-moving in a common direction for almost the entire data period. However, there are frequent intersections, making the two series (to some extent) capable of establishing cointegration in the long-run.

Turning first to the single equation cointegration test results, there are no signs of long-run cointegration for any of the three measures of rer and $\tilde{a}$, as demonstrated by both the EG and ECM tests. For the case of rer_def based model (only), cointegration is being established, though at weak statistical significance. However, the FMOLS and DOLS estimators, tested subsequently, yield insignificant long-run BS coefficients, thus failing to support the presence of a BS effect for Sri Lanka.

Similar results obtain when the models are tested using the multivariate cointegration method. A large majority of the test statistics (Trace and Maximum Eigenvalue) consistently produce a rank of zero. The only exception can be seen under specification 3 of the Johansen ML test where the Eigenvalue statistic produces a rank of 1 for the rer_def based model. Though the evidence is weak in support of cointegration between the model variables, I proceed to estimate a VEC model using the rer_def variable to determine if there is any evidence of a possible BS effect for this real exchange series.

TABLE 5.21: Cointegration Tests Results for Sri Lanka (1981-2010)[30]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[31]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	1	-0.08 [-0.79]	-	-	No
rer_def	No	1	-0.06 [-1.83]	1.73 [1.57]	1.20 [0.57]	No
rer_def_nt	No	1	-0.02 [-1.25]	-	-	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	0	0	No
rer_def	1	0	See below
rer_def_nt	0	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_def	0	Yes	-0.01 [-2.04]	-0.02 [-2.63]	-16.43 [-4.23]	No

[30] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[31] t-values are given in squared-brackets.

The VEC results indicate an inverse BS effect. The first pre-condition, mandatory for the existence of a valid effect, is met, albeit the size of the speed of adjustment coefficient is very small. Movements in the rer series impart a statistically significant correction to short-run fluctuations in rer. But the other two conditions of the model are violated. The condition of weak exogeneity is not met, as can be seen from the second error correction coefficient in TABLE 5.21. And finally, the effect fails to establish because the long-run BS coefficient, though statistically significant, has a negative sign, suggesting a counter-intuitive effect on rer, induced by movements in inter-country sectoral productivities of Sri Lanka and the U.S. Thus, conclusively, the BS effect does not hold for Sri Lanka.

5.5.7 Thailand

A visual inspection of Thailand's rer and $\tilde{a}$ series against the U.S. displays substantial support in favour of a valid long-run relationship. The series are trending in a similar direction throughout the sample period, and there is evidence of the series moving apart and coming together again.

However, contrary to my visual analysis, the single cointegration test results do not provide a lot of support for the BS effect. The tau-values of the CPI and GDP deflator (nontradables) based models when tested against MacKinnon (1996) critical values do not indicate valid long-run causality between the two series. Thus, the country's sectoral productivity gap (against the U.S.) does not appear to significantly cause rer movements in the long-run as measured by these series. The only support is obtained in the case of the GDP deflator based model where ECM (only) yields a negative and significant (though at a weak significance level) EC coefficient. The long-run BS coefficients generated through the FMOLS and DOLS cointegration regression models are positive and significant, indicating the existence of a BS effect for Thailand for this series.

The multivariate cointegration test results do not provide statistical support in favour of the presence of a BS effect. In the case of the GDP deflator (nontradables) rer, the Trace value of the Johansen ML cointegration test (Case 3) produces a valid cointegrating vector. However, the subsequent VEC model does not produce a supporting EC term. Further, there is evidence of endogeneity via the error correction equation for $\tilde{a}$. As a result, I conclude that the evidence from the multivariate analysis is that the BS effect does not hold for Thailand.

5.5.8 Summary of Individual Country Studies

TABLE 5.23 summarizes the preceding results for the country-by-country studies. The right-most column presents an overall summary for each country, based on the individual tests. For the eight countries in the table, I find no support for the BS hypothesis for three of the countries (Indonesia, Japan, and Sri Lanka), and only mixed support for the remaining five. For none of the eight countries do I find support for the BS hypothesis using the multivariate framework. The single equation approach provides, at best, only weak support. For only one country (Malaysia) do the different rer measures produce a consistent finding of evidence in favour of the BS hypothesis. For the other countries, the single equation results differ depending on which variant of the real exchange rate variable is used. The next section pools the data to see if I can obtain a more definitive conclusion regarding the BS hypothesis.

TABLE 5.22: Cointegration Tests Results for Thailand (1971-2013)[32]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	Greater than I(1)	Inconclusive
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[33]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	1	-0.11 [-1.45]	-	-	No
rer_def	No	1	-0.28 [-1.80]	0.38 [3.33]	0.41 [4.45]	Yes
rer_def_nt	No	1	-0.13 [-1.20]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	1	0	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_def_nt	0	Yes	-0.00 [-0.05]	0.48 [2.43]	0.24 [2.77]	No

32 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
33 t-values are given in squared-brackets.

TABLE 5.23: Does Balassa-Samuelson Effect Hold? Summary of Results by Country

Country	Individual Results				Country Summary
	Version of rer	*Single Equation Cointegration Method*	*Multivariate Cointegration Method*		
			Case 3	*Case 4*	
Indonesia (1976-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Japan (1970-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Korea (1970-2013)	rer_cpi	Mixed	No	No	Mixed
	rer_def	Yes	No	No	
	rer_def_nt	No	No	No	
Malaysia (1980-2013)	rer_cpi	Yes	No	No	Mixed
	rer_def	Yes	No	No	
	rer_def_nt	Yes	No	No	
Pakistan (1973-2008)	rer_cpi	No	No	No	Mixed
	rer_def	No	No	No	
	rer_def_nt	Yes	No	No	
Philippines (1971-2013)	rer_cpi	Yes	No	No	Mixed
	rer_def	Yes	No	No	
	rer_def_nt	No	No	No	
Sri Lanka (1981-2010)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	

Country	**Individual Results**				**Country Summary**
	Version of rer	*Single Equation Cointegration Method*	*Multivariate Cointegration Method*		
			Case 3	*Case 4*	
Thailand (1971-2013)	rer_cpi	No	No	No	Mixed
	rer_def	Yes	No	No	
	rer_def_nt	No	No	No	

5.6 Panel Data Estimations

Panel data analysis has a potential advantage over the analysis of individual country data because it allows the pooling of data, providing better statistical power. Most time series suffer from the problem of a small number of observations. This issue results in insignificant t-ratios or F-statistics, raising concerns about the validity and power of short-run as well as long-run estimates. Typically, this issue is common in annual data studies where it is rare to find economic data series covering more than fifty years. In this respect, panel data estimation methods hold an edge because data series can be pooled into panels of different countries. This mitigates the problem of small numbers of observations for researchers.

5.6.1 STEP 0: Panel Unit Root and Stationarity Tests

Up until recently, panel data analysis has not paid much attention to the issues of nonstationarity and cointegration. Nevertheless, owing to interest in relationships among macroeconomic variables, many or most of which are found to be nonstationary, there has been much recent research in this area. Panel unit roots and cointegration tests extend research on univariate time series unit root and cointegration tests. The difference between the two types of tests lies in the asymptotic behaviour of the time series and cross-sectional dimensions.

(i) Combining p-values Test- Fisher-Type Panel Unit Root Tests

The first panel unit root test applied here was developed by Maddala and Wu (1999). Maddala and Wu used the Fisher test to propose a method for combining the p-values from unit root tests from each individual cross-section i to obtain a test statistic for the full panel. The test is nonparametric and has a chi-square distribution with $2n$ degrees of freedom. The test statistic is given by:

(5.11) $\lambda = -2\sum_{i=1}^{n} log_e\,(p_i)$,

where p_i are the p-values from the unit root tests for each cross section i. Note that under the null hypothesis of unit root, $-2log(p_i)$ has a χ^2 distribution with 2 degrees of freedom. For each cross-section, an ADF equation is estimated, and ADF *t*-statistics are computed for each individual series. Then, the corresponding p-values are computed (through Monte Carlo experiments) from the empirical distribution of the ADF test, leading to the appropriate critical values.

The Maddala-Wu test is preferred to other tests for two reasons. Firstly, one is free to apply any unit root test in each time series. The test does not necessitate employing the same test in each cross-section. Secondly, unlike many other panel unit root tests, this test does not require panels to be balanced.

(ii) Hadri Residual-Based LM Stationarity Test

Derived from the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test, Hadri (2000) put forward the Residual based LM test. Based on the OLS residuals obtained from regressing *rer* on a constant and/or time trend, similar to the KPSS test, the Hadri test sets the null hypothesis as the absence of a unit root (stationarity) against the alternative of a unit root in the panel.

(5.12-a) $z_{it} = c_{it} + u_{it}$

(5.12-b) $c_{it} = c_{i,t-1} + v_{it}$

(5.12-c) $H_0 = \sigma_v^2$=0

Where z_{it} = three versions of rer_{it} and $\tilde{a}_{it}$. Assuming $v_{it}$to be zero, this gives rise to a constant c_{it} and, as a consequence, z_{it} will be stationary. The tests allows heteroskedasticity. Provided the number of observations and cross-sections are reasonably large, the empirical size of the test is found to be very close to its nominal size.

Upon finding rer and $\tilde{a}$ first-difference stationary, the next step in panel data analysis is to determine whether a long-run cointegrating relationship exists and to verify that it is consistent with the BS hypothesis. The growing interest in establishing long-run associations in panels has led to the development of various statistical techniques. The most extensively used panel cointegration tests are: Pedroni (1999), Pedroni (2004), Kao (1999) and a Fisher-type test using an underlying Johansen methodology (Maddala and Wu, 1999). The Pedroni and Kao Tests are derived from the Engle-Granger (1987) two-step (residual-based) cointegration tests. The Fisher test is a combined Johansen test. In this analysis, we employ two kinds of panel cointegration tests: Pedroni's (1999), and Johansen's (1988) Fisher panel cointegration tests.

5.6.2 STEP 1: Panel Cointegration Tests

Similar to individual country studies, for establishing a long-run cointegrating relationship between rer and $\tilde{a}$ panels, I shall once again make use of single equation cointegration test and multivariate cointegration approach. Let's discuss each of the two approaches individually.

(i) Pedroni Residual Based Cointegration Test

Developed by Pedroni (1999), the heterogeneous panel cointegration test allows cross-sectional interdependence with individual effects. Provided the data series are unit root in levels, that is, I (1), the Pedroni residual-based cointegration test is an extensively used tool to investigate if a long-run cointegrating association exists between model variables. The following time series panel formulation is proposed by Pedroni:

(5.13-a) $rer_{it} = \alpha_i + \gamma_{it} t + \beta_i \tilde{a}_{it} + \varepsilon_{it}$

(5.13-b) $\hat{\varepsilon}_{it} = \sigma_i \hat{\varepsilon}_{it-1} + \mu_{it}$

Here $i = 1, \dots \dots .., N$ identifies the panels and $t = 1, \dots \dots .., T$ represents time periods. The parameters α_i and $\gamma_{it}t$ are responsible for capturing country-specific effects and deterministic trend effects, respectively. $\hat{\varepsilon}_{it}$ represents the calculated residual deviations from long-run association between rer and $\tilde{a}$. In order to test the null hypothesis of 'No Cointegration' in a panel, that is, $\sigma_i = 1$, Pedroni developed test statistics with asymptotic and finite sample properties. The Pedroni model allows heterogeneity among every member of the panel. Not only this, but the model also allows heterogeneity in long-run cointegrating vectors as well as long-run dynamics.

Under the Pedroni cointegration model, there are actually two sets of residual based tests. The first set of tests consists of pooling the residuals obtained from within-group regressions. The statistics of the tests are standard, normal and asymptotically distributed. This first set of tests includes panel v-statistics, panel ρ-statistics, panel PP-statistics (or t-statistics, non-parametric) and panel ADF-statistics (or t-statistics, parametric). The other group of tests are also standard, normal and asymptotically distributed, but unlike the first set of tests, these tests involve pooling the residuals between the groups. This set consists of group ρ-statistics, group PP-statistics (or t-statistics, non-parametric) and group ADF-statistics (or t-statistics, parametric). All of these seven tests involve estimators that average the estimated coefficients of individual members of the panel. Each of these tests is capable of accommodating individual specific short-run dynamics, individual specific fixed effects and deterministic trends, and individual specific slope coefficients (Pedroni, 2004).

In the event of rejection of the null hypothesis by all seven tests, one may easily draw a conclusion. However, unfortunately, this does not often happen. One frequently confronts a situation where there is a mix of evidence. If this happens, there is a need to look for a test that will explain the power of the cointegration model. As elaborated by Pedroni (2004), in case of a sufficiently large panel, where the issue of size distortion is of little importance, panel v-statistics display the best power in comparison to the other six tests. The panel v-statistics is a one-sided test where the large positive values tend to reject the null hypothesis (Pedroni, 2004). On the other hand, in the case of very small sized panels, group ρ-statistics are likely to reject the null hypothesis. One can be confident enough of the group ρ-statistics as the tests are purposely built for smaller samples and they are regarded as the most conservative of all the seven tests. The rest of the five tests lie somewhere in between the two extreme cases of panel v-statistics and group ρ-statistics. However, they have advantages over a range of large, medium or small sized samples. One noticeable fact is that other than panel v-statistics, the rest of the six tests diverge to negative infinity, that is, the large negative values tend to reject the null hypothesis.

(ii) Fisher-Johansen Combined Panel Cointegration Test

Fisher (1932) derived a combined test that uses the results of individual independent tests. Maddala and Wu (1999) use Fisher's result to propose an alternative approach to testing cointegration in panel data by combining tests from individual cross-sections to obtain a test statistic for the full panel. If p_i is the p-value from an individual cointegration test for cross-section i, then under the null hypothesis for the panel:

(5.14) $-2\sum_{i=1}^{N}\log(p_i) \rightarrow \chi^2(2N)$

Maddala and Wu proposed two statistics: the Fisher statistic from the trace test and the Fisher statistic from the Maximum Eigenvalue test. By default the χ^2 value based on the MacKinnon-Haug-Michelis (1999) p-value is used for Johansen's cointegration Trace test and Maximum Eigenvalue test. Following Johansen's Cointegration approach, cointegration requires the rank to be less than the number of variables in the LR equation.

5.6.3 STEP 2: Estimating the Long-run Relationship between rer and $\tilde{\alpha}$

The second and the final step of the cointegration procedure, serving both the single equation and multivariate cointegration approach, requires the estimation of a long-run BS coefficient. The long-run BS coefficient will be estimated by using the Panel Fully Modified OLS (PFMOLS) and Panel Dynamic OLS (PDOLS) cointegration regression estimators. The two asymptotically unbiased estimators are efficient enough to accommodate considerable heterogeneity across individual members of the panel.

5.6.4 Results

I start my panel data estimations by formally testing the three model variables using panel unit root tests. I test the variables using Fisher-ADF and Fisher-PP unit root tests and the Hadri stationarity test.

The results are reported in the first panel of TABLE 5.24. For all three versions of rer, the three tests generate common results. The three rer measures are a unit root process in levels according to the Fisher-ADF test, Fisher-PP test and Hadri test. rer turns out to be integrated of order one at the one percent significance level. This is true for all three variants of rer except rer_cpi, for which the Fisher-ADF unit root test reaches the conclusion that the series is I(0). However, following the majority of empirical evidence, I conclude the series to be a unit root process in levels.

For the $\tilde{\alpha}$ series, the two unit root tests indicate it to be nonstationary in levels. However, the Hadri stationarity test reveals that the order of integration of $\tilde{\alpha}$ is greater than I(1). Keeping in view the majority of empirical evidence, I conclude the series to be a unit root process in levels.

The second and third panels of TABLE 5.24 display the test results for the Pedroni residual based cointegration test and the Johansen Fisher panel cointegration test. Discussing the test statistics obtained from the Pedroni cointegration test first, I opted for automatic lag selection. All the seven test statistics unanimously failed to reject the null hypothesis of no cointegration for the cointegration models using rer_def and rer_def_nt as regressands. The only exception is the rer_cpi based long-run model. However, for this version as well, the dominant number of test statistics (5 out of 7 tests) report evidence in support of no cointegration between model variables. Given that 5 of the 7 tests fail to reject the null of cointegration, including the Panel *v* and Group *rho* tests, I interpret these results as supporting the inexistence of the BS effect for the cross-sectional data set of Asian economies.

As regarding the test results obtained from the Fisher-Johansen panel cointegration test, similar to the Pedroni cointegration test, the results are once again not supportive of a valid long-run association between the model variables. As the test requires the user to specify lag lengths, I selected the lag length through Panel VAR, following the lag suggestion of SIC. The two individual test specifications yield different results. The Trace and Maximum Eigenvalue statistics of specification 3 of the test found no evidence of a valid cointegrating vector for all three estimated models. The test statistics commonly produce a rank of 2, challenging my unit root test findings, proving the model variables to be level stationary. However, specification 4 of the tests provided some support for a valid long-run association for the rer_cpi based real exchange rate (though marginally) and the rer_def_nt based long-run models. However, in the former case, this evidence was insufficient to proceed further because two out of the three cointegration tests (Pedroni cointegration and Trace statistics) indicated the absence of long-run co-movement between model variables. With respect to rer_def_nt, I proceeded by using PFMOLS and PDOLS to estimate the long-run exchange rate equation, but the BS coefficient was statistically insignificant.

TABLE 5.25 summarizes the results of the panel analysis. For all three versions of the real exchange rate variable, I find no evidence in favour of the BS hypothesis.

Putting it all together, I conclude that there is little evidence in support of a BS effect for the emerging Asian countries. There is some, generally weak and mixed support, for the BS hypothesis when using the individual country data. However, only for the single equation models, and generally not for all variants of the real exchange rate variable. The multivariate models find no support for the BS hypothesis using the individual country data. Likewise, there is no support for the BS hypothesis when the data are pooled. Perhaps the reason for this lack of success is due to how industries are classified as tradable or nontradable. I pursue this line of analysis in the next chapter.

TABLE 5.24: Summary of Results for Panel Unit Root and Cointegration Tests for Balassa-Samuelson Effect[34,35]

Panel Unit Root and Stationarity Test Results (Order of Integration as Determined by)

Variables	Fisher-ADF	Fisher-PP	Hadri	Conclusion
rer_cpi	I (0)**	I (1)***	I (1)***	I (1)
rer_def	I (1)***	I (1)***	I (1)***	I (1)
rer_def_nt	I (1)***	I (1)***	I (1)***	I (1)
$\tilde{a}$	I (1)***	I (1)***	Greater than I (1)	I (1)

Pedroni Panel Cointegration Test Results[36]

	Common AR Coefficients (Within Dimension)				*Individual AR Coefficients (Between Dimension)*			
Version of rer	Panel v Statistics	Panel ρ Statistics	Panel PP Statistics	Panel ADF Statistics	Group ρ Statistics	Group PP Statistics	Group ADF Statistics	Does BS Effect Hold?
rer_cpi	0.23	-0.37	-0.56	-1.32*	0.70	0.05	-2.38***	No
rer_def	-0.54	0.68	0.71	-0.06	1.47	1.13	-0.31	No
rer_def_nt	-0.69	0.52	0.27	0.20	1.23	0.57	-0.68	No

34 ***, ** and * are representing significance of sample statistics at 1%, 5% and 10% levels respectively.
35 Hong Kong is omitted from panel estimations.
36 Pedroni panel cointegration is a test for null of no cointegration in both homogenous and heterogeneous panels. The test statistics are standardized and asymptotically normally distributed. See Pedroni (1995, 1999) for further details.

Johansen-Fisher Panel Cointegration Test Results[37,38]

Case 3: Intercept (no trend) in cointegrating equation and VAR

Version of rer	Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigen Stat)	Does BS Effect Hold?
rer_cpi	2	2	No
rer_def	2	2	No
rer_def_nt	2	2	No

Case 4: Intercept and trend in cointegrating equation-no trend in VAR

Version of rer	Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigen Stat)	Does BS Effect Hold?
rer_cpi	1	0	N/A[39]
rer_def	0	0	No
rer_def_nt	1	1	See below

Results for Panel FMOLS and DOLS[40] **Estimators**
Long-run Cointegrating Vectors for Balassa-Samuelson Effect

Version of rer	**Estimator**	**BS Coefficient**[41]	**Does BS Effect Hold?**
rer_def_nt	PFMOLS	0.04 [0.74]	No
	PDOLS	0.09 [1.59]	No

[37] The test is maximum likelihood based rank test.
[38] Lag selection is done through SIC under panel VAR.
[39] Insufficient evidence in support of cointegration between rer_cpi and $\tilde{a}$ (P1.A and P1.C both support inexistence of cointegration) prevents me from proceeding further with PFMOLS and PDOLS estimations.
[40] Lead = Lag = 1.
[41] t-values are given in squared-brackets.

TABLE 5.25: Does Balassa-Samuelson Effect Hold?
Summary of Results for Panel Cointegration Tests

Version of rer	Tests of Cointegration: Pedroni Residual Based Panel Cointegration Test	Johansen-Fisher Panel Cointegration Test: Case 3	Johansen-Fisher Panel Cointegration Test: Case 4	Conclusion
rer_cpi	No	No	No	No
rer_def	No	No	No	No
rer_def_nt	No	No	No	No

APPENDIX-A

EViews Programming Code for Korea

```
wfopen "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-5\korea.wf1"

'*****************************************************************************
'Group Plot for RER_CPI, RER_DEF, RER_DEF_NT and A_TILDE
'*****************************************************************************
group gA rer_cpi rer_def rer_def_nt A_tilde
freeze(group_plot) gA.line(x)
group_plot.setelem(1) lcolor(black) symbol(7) lpat(1)
group_plot.setelem(2) lcolor(black) symbol(4) lpat(1)
group_plot.setelem(3) lcolor(black) symbol(1) lpat(1)
group_plot.setelem(3) lcolor(black)
group_plot.options linepat
group_plot.addtext(t) Real ERs and Productivity Gap (Korea & U.S): 1970-2013
group_plot.addtext(b) Year
group_plot.addtext(l) RER_CPI
group_plot.addtext(l) RER_DEF
group_plot.addtext(l) RER_CPI, RER_DEF & RER_DEF_NT
group_plot.addtext(r) A_TILDE
'**************************************************************
'**************************************************************
create y 1970 2013
'importing data from Excel for Korea
import "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-5\Chapter 5.xlsx" range="Korea"

'*****************************************************************************************************
'CASE-1: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_CPI & A_TILDE
'*****************************************************************************************************
'*****************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'*****************************************************************
'**********************************
'Graph for Korea's RER_CPI
'**********************************

genr rer_cpi = rer_cpi
freeze(figure_rer_cpi) rer_cpi.line
figure_rer_cpi.addtext(t) rer_cpi (Korea):  1970-2013
figure_rer_cpi.addtext(b) Year
figure_rer_cpi.addtext(l) rer_cpi
figure_rer_cpi.legend(off)

'We see from the FIGURE that rer_cpi has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'************************************************
'ADF Unit Root Test for Korea's RER_CPI
'************************************************

freeze(table_5_3_rer_cpi_adf) rer_cpi.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 1.  The unit root test produces a t-value of -3.32 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.
```

```
genr resid = 0
freeze(mode=overwrite,rer_cpi_adf) rer_cpi.uroot(adf,const,trend,info=sic)
freeze(rer_cpi_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_cpi series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr rer_cpidiff = d(rer_cpi)
freeze(figure_rer_cpidiff) rer_cpidiff.line
figure_rer_cpidiff.addtext(t) Drer_cpi (Korea):  1970-2013
figure_rer_cpidiff.addtext(b) Year
figure_rer_cpidiff.addtext(l) drer_cpi
figure_rer_cpidiff.legend(off)
```

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

```
genr rer_cpidiff = d(rer_cpi)
freeze(table_5_4_rer_cpidiff1_adf) rer_cpidiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -5.42 which is now smaller than our 5% criterion -2.93. Thus, we may now reject the null of non-stationarity in first differenced series of rer_cpi. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_cpidiff1_adf) rer_cpidiff.uroot(adf,const,info=sic)
freeze(rer_cpidiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_cpi series is I(1).

'***
'DF-GLS Unit Root Test for Korea's RER_CPI
'***

```
freeze(table_5_5_rer_cpi_dfgls) rer_cpi.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -3.39 which is smaller than our 5% criterion -3.19. Thus, at this point, we may reject the null of unit root.

''Putting it all together, I conclude that the rer_cpi series is I(0), a finding incompatible with my ADF test results.

'**
'Graph for Korea's Productivity (a_tilde)
'**

```
genr a_tilde = a_tilde
freeze(figurea_tilde) a_tilde.line
figurea_tilde.addtext(t) a_tilde (Korea):  1970-2013
figurea_tilde.addtext(b) Year
figurea_tilde.addtext(l) a_tilde
figurea_tilde.legend(off)
```

'We see from the FIGURE that a_tilde has time trend to it. So we would include both an intercept and a time trend in our unit root regression equations.

```
'***************************************************
'ADF Unit Root Test for Korea's Productivity
'***************************************************

freeze(table_5_3_a_tilde_adf) a_tilde.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.42 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,a_tilde_adf) a_tilde.uroot(adf,const,trend,info=sic)
freeze(a_tilde_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the a_tilde series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr a_tildediff = d(a_tilde)
freeze(figure_a_tildediff) a_tildediff.line
figure_a_tildediff.addtext(t) da_tilde (Korea):  1970-2013
figure_a_tildediff.addtext(b) Year
figure_a_tildediff.addtext(l) da_tilde
figure_a_tildediff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr a_tildediff = d(a_tilde)
freeze(table_5_4_a_tildediff1_adf) a_tildediff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0.  The unit root test produces a t-value of -6.51 which is now smaller than our 5% criterion -2.93.  Thus, we may now reject the null of non-stationarity in first differenced series of a_tilde.  There is no reason to go further.  The last thing we do is to check ADF regression result for white noise.

genr resid = 0
freeze(mode=overwrite,a_tildediff1_adf) a_tildediff.uroot(adf,const,info=sic)
freeze(a_tildediff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the a_tilde series is I(1).

'*******************************************************
'DF-GLS Unit Root Test for Korea's Productivity
'*******************************************************

freeze(table_5_5_a_tilde_dfgls) a_tilde.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.48 which is greater than our 5% criterion -3.19.  Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr a_tildediff = d(a_tilde)
freeze(table_5_6_a_tildediff1_dfgls) a_tildediff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 0.  The unit root test produces a t-value of -6.53 which is now smaller than our 5% criterion -1.94. Thus, we may reject the null of non-stationarity in first differenced series of a_tilde.

''Putting it all together, I conclude that the a_tilde series is I(1), a finding compatible with my ADF test results.
```

```
'************************************************
'Single Equation Cointegration Methods
'************************************************
'**************************************************************
''Graph the suspected cointegrated series together
'**************************************************************

'The first step is to plot a graph of the suspected series.  This is very important!

group g1 rer_cpi a_tilde
freeze(figure5_3a) g1.line(x)
figure5_3a.setelem(1) lcolor(black)
figure5_3a.setelem(2) lcolor(black) lpat(8)
figure5_3a.options linepat
figure5_3a.addtext(t) rer_cpi and a_tilde (Korea & U.S): 1970-2013
figure5_3a.addtext(b) Year
figure5_3a.addtext(l) rer_cpi
figure5_3a.addtext(r) a_tilde

''***********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'***********************************************************

freeze(table_5_8_egc_rer_cpi) g1.coint(method=eg)

'The null hypothesis will be rejected as suggested by sample statistics.

''*********************************************
''S1.B.Error Correction Model (ECM)
'*********************************************

''*************************************************
'Selecting the number of lags in the VAR
'*************************************************

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var var1.ls 1 4   g1
freeze(var1_lagtest1) var1.laglen(4)
freeze(var1_lagtest2) var1.testlags

'The lag length test above indicates that the VAR has 1 lags. But the residuals are not absolutely white noise. So I raised the number of lags from 1 to 4 but still did not obtain white residuals. So, I continue with my actual lag specification of 1.

var var2.ls 1 1  g1
freeze(var2_artest1) var2.correl
freeze(var2_artest2) var2.qstats(12)
freeze(var2_artest3) var2.arlm(12)

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_cpi c a_tilde
genr ec1 = resid


var table_5_10_eg2a_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1))

var table_5_10_eg2b_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1)) d(a_tilde(-1))

var table_5_10_eg2c_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.

var table_5_10_eg2a_cpi.ls 0 0 d(rer_cpi)   @   c ec1(-1) d(rer_cpi(-1))
```

```
freeze(table_5_10_eg2a1_cpi_artest1) table_5_10_eg2a_cpi.correl
freeze(table_5_10_eg2a2_cpi_artest2) table_5_10_eg2a_cpi.qstats(12)
freeze(table_5_10_eg2a3_cpi_artest3) table_5_10_eg2a_cpi.arlm(12)
'The residuals are not absolutely white noise. Let's try other two models.

var table_5_10_eg2b_cpi.ls 0 0 d(rer_cpi)   @   c ec1(-1) d(rer_cpi(-1)) d(a_tilde(-1))
freeze(table_5_10_eg2b1_cpi_artest1) table_5_10_eg2b_cpi.correl
freeze(table_5_10_eg2b2_cpi_artest2) table_5_10_eg2b_cpi.qstats(12)
freeze(table_5_10_eg2b3_cpi_artest3) table_5_10_eg2b_cpi.arlm(12)

'The residuals are still not white noise.

var table_5_10_eg2c_cpi.ls 0 0 d(rer_cpi)   @   c ec1(-1) d(rer_cpi(-1)) d(a_tilde(-1)) d(a_tilde(-2))
freeze(table_5_10_eg2c1_cpi_artest1) table_5_10_eg2c_cpi.correl
freeze(table_5_10_eg2c2_cpi_artest2) table_5_10_eg2c_cpi.qstats(12)
freeze(table_5_10_eg2c3_cpi_artest3) table_5_10_eg2c_cpi.arlm(12)

'None of the above three models generate white residuals. This makes me to raise the number of lags of rer_cpi from 1 to 3. Starting with Model A, I am going to raise the number of lags of RER_CPI to 3.

var eg2d_cpi.ls 0 0 d(rer_cpi)   @   c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3))
freeze(table_5_10_eg2d1_cpi_artest1) eg2d_cpi.correl
freeze(table_5_10_eg2d2_cpi_artest2) eg2d_cpi.qstats(12)
freeze(table_5_10_eg2d3_cpi_artest3) eg2d_cpi.arlm(12)

'But even by adding 3 lags of rer_cpi, we do not obtain white residuals. So, let's continue with our very first model i.e. Model A and estimate EC model.

''*************************
'Estimating EC Model
'**************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_5_11_ecm_rer_cpi.ls(n) d(rer_cpi) c ec1(-1) d(rer_cpi(-1))

'Note that the SR effect is significant as the error correction coefficient -0.38 is statistically significant at better than 1% significance level.

''***********************************************
''S2.A & S2.B: Obtaining LR Coefficients
'************************************************

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Korea against U.S.

equation table_5_12_LReqn1a_fmols.cointreg(method=fmols) rer_cpi a_tilde

equation table_5_12_LReqn1b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2 ) rer_cpi a_tilde

'The BS coefficient obtained through FMOLS estimator is 0.13 and is statistically insignificant whereas the one generated through DOLS test 0.16 and is statistically significant at 10% significance level. Thus, there is 'Mixed' evidence in support of BS effect existing for Korea.

''*******************************************
'Multivariate Cointegration Approach
'*******************************************

''*********************************************************
''Check if the VAR (2) model is dynamically stable
'**********************************************************
freeze(var2_varstable) var2.arroots(graph)

'The model is dynamically stable.
```

''**
''M1.A & M1.B: Identifying the number of cointegrating vectors
'**

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

```
freeze(var2_table_5_14_coint1) var2.coint(s,1)
```

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. All the results indicate 0 cointegrating vectors.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

'CONCLUSION: I conclude that rer_cpi and a_tilde are not cointegrated in the Korean data.

'***
'CASE-2: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_DEF & A_TILDE
'***
'**
'STEP 0: Tests for Unit Root in Individual Time Series
'**
'**********************************
'Graph for Korea's RER_DEF
'**********************************

```
genr rer_def = rer_def
freeze(figure_rer_def) rer_def.line
figure_rer_def.addtext(t) rer_def (Korea):  1970-2013
figure_rer_def.addtext(b) Year
figure_rer_def.addtext(l) rer_def
figure_rer_def.legend(off)
```

'We see from the FIGURE that rer_def has time trend to it. So we would include both an intercept and a time trend in our unit root regression equations.

'**
'ADF Unit Root Test for Korea's RER_DEF
'**

```
freeze(table_5_3_rer_def_adf) rer_def.uroot(adf,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -3.35 which is greater than our 5% criterion -3.52. Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_def_adf) rer_def.uroot(adf,const,trend,info=sic)
freeze(rer_def_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr rer_defdiff = d(rer_def)
freeze(figure_rer_defdiff) rer_defdiff.line
figure_rer_defdiff.addtext(t) drer_def (Korea):  1970-2013
```

```
figure_rer_defdiff.addtext(b) Year
figure_rer_defdiff.addtext(l) drer_def
figure_rer_defdiff.legend(off)
```

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

```
genr rer_defdiff = d(rer_def)
freeze(table_5_4_rer_defdiff1_adf) rer_defdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -5.49 which is now smaller than our 5% criterion -2.93. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_defdiff1_adf) rer_defdiff.uroot(adf,const,info=sic)
freeze(rer_defdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def series is I(1).

```
'*******************************************************
'DF-GLS Unit Root Test for Korea's RER_DEF
'*******************************************************
```

```
freeze(table_5_5_rer_def_dfgls) rer_def.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -3.43 which is smaller than our 5% criterion -3.19. Thus, at this point, we may reject the null of a unit root. These findings are in contrast with ADF test results. Thus, rer_def series is I(0) according to DF-GLS test results.

```
'***********************************************
'Single Equation Cointegration Methods
'***********************************************
```

```
'************************************************************
''Graph the suspected cointegrated series together
'************************************************************
```

'The first step is to plot a graph of the suspected series. This is very important!

```
group g2 rer_def a_tilde
freeze(figure5_3b) g2.line(x)
figure5_3b.setelem(1) lcolor(black)
figure5_3b.setelem(2) lcolor(black) lpat(8)
figure5_3b.options linepat
figure5_3b.addtext(t) rer_def and a_tilde (Korea & U.S): 1970-2013
figure5_3b.addtext(b) Year
figure5_3b.addtext(l) rer_def
figure5_3b.addtext(r) a_tilde
```

```
''*********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'*********************************************************
```

```
freeze(table_5_8_egc_rer_def) g2.coint(method=eg)
```

'The null hypothesis will not be rejected as suggested by sample statistics.

```
''********************************************
''S1.B.Error Correction Model (ECM)
'********************************************
```

```
''**************************************************
'Selecting the number of lags in the VAR
'**************************************************

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.
var var3.ls 1 4   g2
freeze(var3_lagtest1) var3.laglen(4)
freeze(var3_lagtest2) var3.testlags

'The lag length test above indicates that the VAR has 1 lags. But the residuals are not absolutely white noise. So I am not satisfied with the selection of 1 lag. I raised the number of lags from 1 to 4 and thus obtained somewhat white residuals.

var var4.ls 1 4  g2
freeze(var4_artest1) var4.correl
freeze(var4_artest2) var4.qstats(12)
freeze(var4_artest3) var4.arlm(12)

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_def c a_tilde
genr ec2 = resid

var table_5_10_eg2a_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4))

var table_5_10_eg2b_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4)) d(a_tilde(-1))

var table_5_10_eg2c_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.

var table_5_10_eg2a_def.ls 0 0 d(rer_def)  @   c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4))
freeze(table_5_10_eg2a1_def_artest1) table_5_10_eg2a_def.correl
freeze(table_5_10_eg2a2_def_artest2) table_5_10_eg2a_def.qstats(12)
freeze(table_5_10_eg2a3_def_artest3) table_5_10_eg2a_def.arlm(12)

'The residuals are not absolutely white noise. Let's try other two models.

var table_5_10_eg2b_def.ls 0 0 d(rer_def)   @   c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4)) d(a_tilde(-1))
freeze(table_5_10_eg2b1_def_artest1) table_5_10_eg2b_def.correl
freeze(table_5_10_eg2b2_def_artest2) table_5_10_eg2b_def.qstats(12)
freeze(table_5_10_eg2b3_def_artest3) table_5_10_eg2b_def.arlm(12)

'The residuals are still not white noise.

var table_5_10_eg2c_def.ls 0 0 d(rer_def)   @   c ec2(-1) d(rer_def(-1)) d(a_tilde(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4)) d(a_tilde(-2))
freeze(table_5_10_eg2c1_def_artest1) table_5_10_eg2c_def.correl
freeze(table_5_10_eg2c2_def_artest2) table_5_10_eg2c_def.qstats(12)
freeze(table_5_10_eg2c3_def_artest3) table_5_10_eg2c_def.arlm(12)

'None of the above three models generate white residuals. So, let's continue with our very first model i.e. Model A and estimate EC model.

''*************************
'Estimating EC Model
'**************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:
```

```
equation table_5_11_ecm_rer_def.ls(n) d(rer_def) c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(rer_def(-4))

'Note that the SR effect is significant as the error correction coefficient -0.99 is statistically significant at better than 1% significance level.

''***********************************************
''S2.A & S2.B: Obtaining LR Coefficients
'************************************************

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Korea against U.S.

equation table_5_12_LReqn2a_fmols.cointreg(method=fmols) rer_def a_tilde

equation table_5_12_LReqn2b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2 ) rer_def a_tilde

'The BS coefficients obtained through FMOLS and DOLS estimators are 0.51 and 0.57 and are statistically significant. Thus, there is sufficient evidence in support of BS effect existing for Korea.

''*******************************************
'Multivariate Cointegration Approach
'*******************************************

''**************************************************************
''Check if the VAR (4) model is dynamically stable
'**************************************************************
freeze(var4_varstable) var4.arroots(graph)

'The model is dynamically stable.

''************************************************************************
''M1.A & M1.B: Identifying the number of cointegrating vectors
'************************************************************************

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(var4_table_5_14_coint2) var4.coint(s,4)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. Trace statistic of Case 3 indicate 1 cointegrating vector.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).


'CONCLUSION:  I conclude that rer_def and a_tilde for Korea and U.S. are cointegrated under single equation cointegration models but are not cointegrated as proven by multivariate cointegration approach.

'*********************************************************************************************************
'CASE-3: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_DEF_NT & A_TILDE
'*********************************************************************************************************
'**************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'**************************************************************
'****************************************
'Graph for Korea's RER_DEF_NT
'**************************************
```

```
genr rer_def_nt = rer_def_nt
freeze(figure_rer_def_nt) rer_def_nt.line
figure_rer_def_nt.addtext(t) rer_def_nt (Korea):  1970-2013
figure_rer_def_nt.addtext(b) Year
figure_rer_def_nt.addtext(l) rer_def_nt
figure_rer_def_nt.legend(off)

'We see from the FIGURE that rer_def_nt has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'**************************************************
'ADF Unit Root Test for Korea's RER_DEF_NT
'**************************************************
 freeze(table_5_3_rer_def_nt_adf) rer_def_nt.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -2.83 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_nt_adf) rer_def_nt.uroot(adf,const,trend,info=sic)
freeze(rer_def_nt_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the rer_def_nt series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr rer_def_ntdiff = d(rer_def_nt)
freeze(figure_rer_def_ntdiff) rer_def_ntdiff.line
figure_rer_def_ntdiff.addtext(t) drer_def_nt (Korea):  1970-2013
figure_rer_def_ntdiff.addtext(b) Year
figure_rer_def_ntdiff.addtext(l) drer_def_nt
figure_rer_def_ntdiff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_5_4_rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0.  The unit root test produces a t-value of -5.16 which is now smaller than our 5% criterion -2.93.  Thus, we may now reject the null of non-stationarity in first differenced series of rer_def_nt.  There is no reason to go further.  The last thing we do is to check ADF regression result for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)
freeze(rer_def_ntdiff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_nt series is I(1).

'******************************************************
'DF-GLS Unit Root Test for Korea's RER_DEF_NT
'******************************************************

freeze(table_5_5_rer_def_nt_dfgls) rer_def_nt.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -2.51 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not
```

```
Genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_5_6_rer_def_ntdiff1_dfgls) rer_def_ntdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -4.93 which is now smaller than our 5% criterion -1.94. Thus, we may reject the null of non-stationarity in first differenced series of rer_def_nt.

''Putting it all together, I conclude that the rer_def_nt series is I(1), a finding compatible with my ADF test results.

```
'**********************************************
'Single Equation Cointegration Methods
'**********************************************

'***********************************************************
''Graph the suspected cointegrated series together
'***********************************************************
```

'The first step is to plot a graph of the suspected series. This is very important!

```
group g3 rer_def_nt a_tilde
freeze(figure5_3) g3.line(x)
figure5_3.setelem(1) lcolor(black)
figure5_3.setelem(2) lcolor(black) lpat(8)
figure5_3.options linepat
figure5_3.addtext(t) rer_def_nt and a_tilde (Korea & U.S): 1970-2013
figure5_3.addtext(b) Year
figure5_3.addtext(l) rer_def_nt
figure5_3.addtext(r) a_tilde
```

```
''********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'********************************************************
```

```
freeze(table_5_8_egc_rer_def_nt) g3.coint(method=eg)
```

'The null hypothesis will not be rejected as suggested by sample statistics.

```
''*******************************************
''S1.B.Error Correction Model (ECM)

''*******************************************
''***********************************************
'Selecting the number of lags in the VAR
'************************************************
```

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

```
var var5.ls 1 4   g3
freeze(var5_lagtest1) var5.laglen(4)
freeze(var5_lagtest2) var5.testlags
```

'The lag length test above indicates that the VAR has 1 lags. But the residuals are not absolutely white noise. So I am not satisfied with the selection of 1 lag. I raised the number of lags from 1 to 4 and thus obtained somewhat white residuals.

```
var var6.ls 1 4  g3
freeze(var6_artest1) var6.correl
freeze(var6_artest2) var6.qstats(12)
freeze(var6_artest3) var6.arlm(12)
```

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

```
genr resid = 0
equation eg.ls rer_def_nt c a_tilde
genr ec3 = resid
```

var table_5_10_eg2a_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4))

var table_5_10_eg2b_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4)) d(a_tilde(-1))

var table_5_10_eg2c_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best. Now we test that model for serial correlation.

var table_5_10_eg2a_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4))
freeze(table_5_10_eg2a1_def_nt_artest1) table_5_10_eg2a_def_nt.correl
freeze(table_5_10_eg2a2_def_nt_artest2) table_5_10_eg2a_def_nt.qstats(12)
freeze(table_5_10_eg2a3_def_nt_artest3) table_5_10_eg2a_def_nt.arlm(12)

'The residuals are not absolutely white noise. Let's try other two models.

var table_5_10_eg2b_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4)) d(a_tilde(-1))
freeze(table_5_10_eg2b1_def_nt_artest1) table_5_10_eg2b_def_nt.correl
freeze(table_5_10_eg2b2_def_nt_artest2) table_5_10_eg2b_def_nt.qstats(12)
freeze(table_5_10_eg2b3_def_nt_artest3) table_5_10_eg2b_def_nt.arlm(12)

'The residuals are still not white noise.

var table_5_10_eg2c_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4)) d(a_tilde(-1)) d(a_tilde(-2))
freeze(table_5_10_eg2c1_def_nt_artest1) table_5_10_eg2c_def_nt.correl
freeze(table_5_10_eg2c2_def_nt_artest2) table_5_10_eg2c_def_nt.qstats(12)
freeze(table_5_10_eg2c3_def_nt_artest3) table_5_10_eg2c_def_nt.arlm(12)

'None of the above three models generate white residuals. So, let's continue with our very first model i.e. Model A and estimate EC model.

''************************
'Estimating EC Model
'*************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_5_11_ecm_rer_def_nt.ls(n) d(rer_def_nt) c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(rer_def_nt(-4))

'Note that the SR effect is significant as the error correction coefficient -0.63 is statistically significant at better than 1% significance level.

''***
''S2.A & S2.B: Obtaining LR Coefficients
'**

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Korea against U.S.

equation table_5_12_LReqn3a_fmols.cointreg(method=fmols) rer_def_nt a_tilde

equation table_5_12_LReqn3b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2) rer_def_nt a_tilde

'The BS coefficients obtained through FMOLS and DOLS estimators are -0.11 and -0.14 and are statistically insignificant. Thus, there is insufficient evidence in support of BS effect existing for Korea.

```
"*********************************************
'Multivariate Cointegration Approach
'********************************************

"************************************************************
"Check if the VAR (6) model is dynamically stable
'*************************************************************
freeze(var6_varstable) var6.arroots(graph)

'The model is dynamically stable.

"***************************************************************************
"M1.A & M1.B: Identifying the number of cointegrating vectors
'****************************************************************************

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(var6_table_5_14_coint3) var6.coint(s,4)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. Trace statistic of Case 3 indicate 1 cointegrating vector.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).


'CONCLUSION:  I conclude that rer_def_nt and a_tilde for Korea and U.S. are cointegrated under single equation cointegration models but are not cointegrated as proven by multivariate cointegration approach.
```

CHAPTER 6: INVESTIGATING THE STANDARD VERSION OF THE BALASSA-SAMUELSON HYPOTHESIS WITH BROAD SECTORAL DIVISION

6.1 Motivation

In Chapter Five, we empirically tested the Balassa-Samuelson (BS) hypothesis for the developing ASEAN and SAARC countries, but did not find much evidence in its favor. These findings are in line with a few existing studies on Asia (Drine and Rault, 2004; Gente, 2006; Wang et al., 2016), but in contrast with a sizeable number of other studies that find evidence in support of the BS effect for the region (Chinn, 2000; Bahmani-Oskooee and Nasir, 2004; Choudhri and Khan, 2005; Thomas and King, 2008; Olson, 2009; Tsen, 2011; Chowdhury, 2012, etc.). The empirical literature lists multiple reasons for why different studies can reach different conclusions. Price and productivity measures may differ in their ability to capture the proposed effect (see Canzoneri et al., 1999; Egert et al., 2003; Lee and Tang, 2007), different econometric procedures may differ in their ability to measure short- and long-run dynamics (see Chong et al., 2012; Boreo et al., 2015), and different groupings of industrial sectors may differ in their ability to correctly categorize the open and sheltered sectors of the economy (see Egert, 2002, 2005; Mihaljek and Klau, 2004; Gibson, 2008). In short, there are several potential reasons behind the failure of the empirical verification of the BS hypothesis.

The former two reasons (as stated above) were addressed in the preceding chapter, employing three alternative measures of the real exchange rate and using different cointegration models for testing the hypothesis. The primary objective of this chapter is to address the latter issue, i.e., to examine the robustness of the results in Chapter Five when the model is tested using more disaggregated sectoral level data. This should give better coverage of industries under traded and nontraded sectors. This will serve as an additional robustness test on the BS hypothesis. It will reveal how sensitive the model results are to finer sectoral categorization.

The estimations are run using the same model/theoretical specifications, real exchange rate (rer) and productivity ($\tilde{a}$) measures, and econometric procedures that were used in Chapter Five. The only difference lies with the sectoral classification, which is finer and broader this time. The real sector of each country is consistently divided into seven distinct sectors. rer and $\tilde{a}$ measures are constructed by categorizing each of these sectors (traded and nontraded), following the sectoral division suggested by Dumrongrittikul (2012). The task was never easy as I confronted multi-faceted issues while organizing output and employment data series in a meaningful way. For sectoral division, the detailed discussions on the procedures involved can be seen in Chapter Three of the dissertation.

The rest of the chapter follows the pattern of Chapter Five. For verifying BS hypothesis empirically with more disaggregated sectoral level data, at first country by country analysis will be done using single equation and multivariate cointegration models. Later, pooled data examination will also be conducted using single equation residual based and multivariate rank cointegration methods to make it easier to quantitatively summarize the overall findings.

6.2 Country Studies

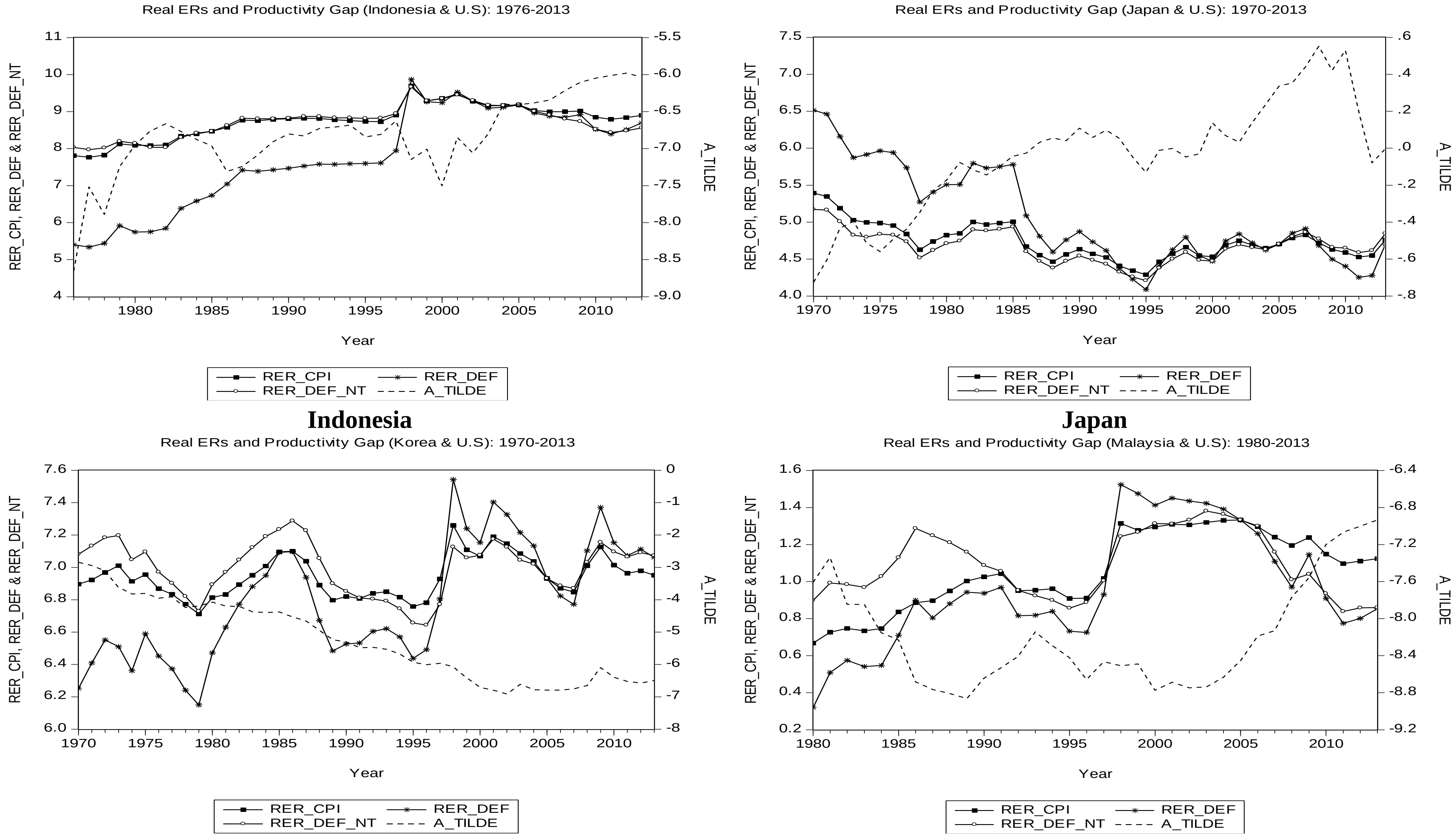
In this section, robust country-by-country results of the BS hypothesis with disaggregated data are reported. In Chapter Five, the country study on Korea provided a detailed report of the various steps involved in testing the BS hypothesis. The EViews programs for Indonesia, which are attached in the appendix to this chapter, provide details of the procedures consistently followed in this chapter when analyzing individual countries.

6.2.1 Indonesia

We begin by displaying the time plots of real exchange rate and cross-country sectoral productivity differentials of Indonesia against the U.S. for the period year 1976 to 2013. In FIGURE 6.1, the three successive plots display CPI, GDP deflator and GDP deflator (nontradables) based rer of Indonesia against the U.S. The solid lines represent the three alternative versions of rer and the dotted line represents $\tilde{a}$. From visual inspection, I obtain considerable amount of support in favor of BS hypothesis. For all three types of rer, the model variables are trending in a similar direction. Moreover, their reasonably frequent intersection makes their long-run association fairly plausible.

Next, we proceed to formally test the rer and $\tilde{a}$ series using two different cointegration tests for estimating the long-run BS effect. Single equation cointegration methods (under Engle-Granger (EG) test specification and error correction representation) and multivariate cointegration methods (under the Johansen ML cointegration procedure) are used to test whether a long term cointegrating relationship exists. The results of the tests are reported in TABLE 6.1 below.

FIGURE 6.1: Plots for Real Exchange Rates and Sectoral Productivity Differentials against U.S.

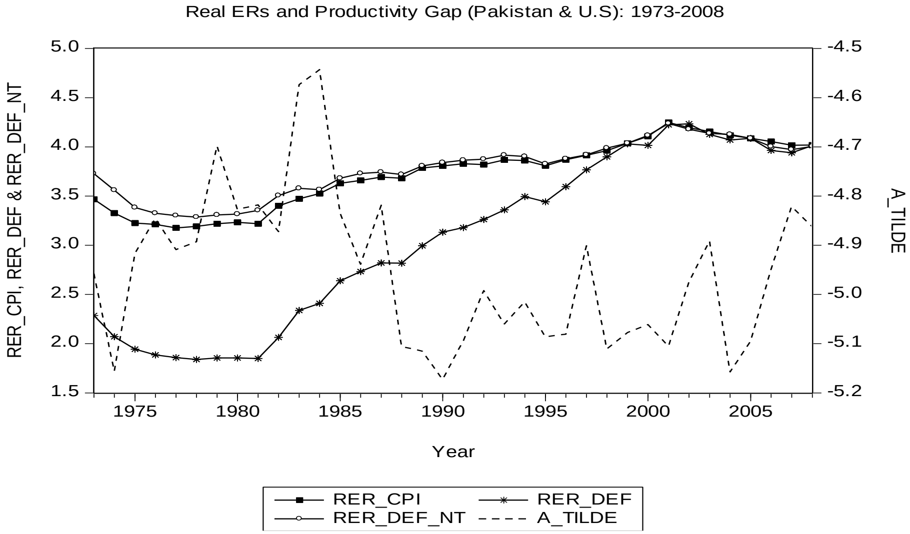


Indonesia

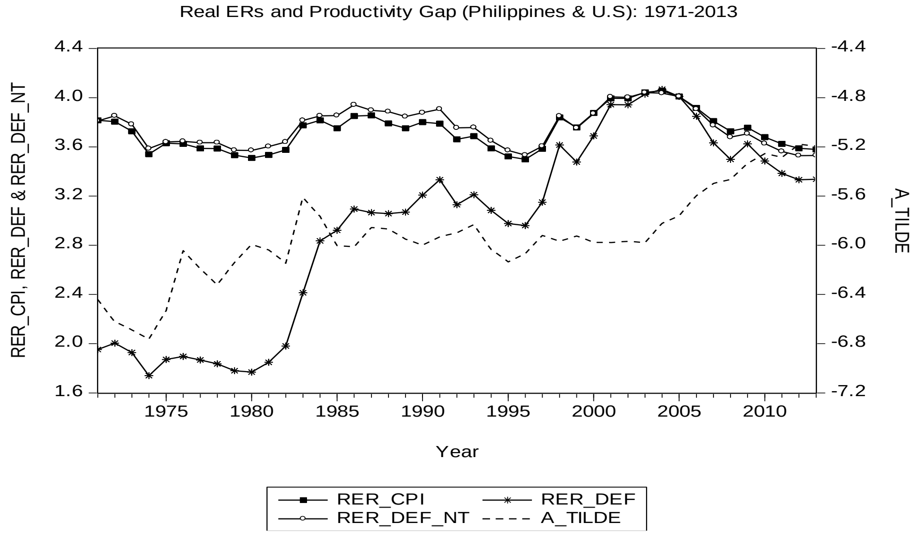
Japan

Korea

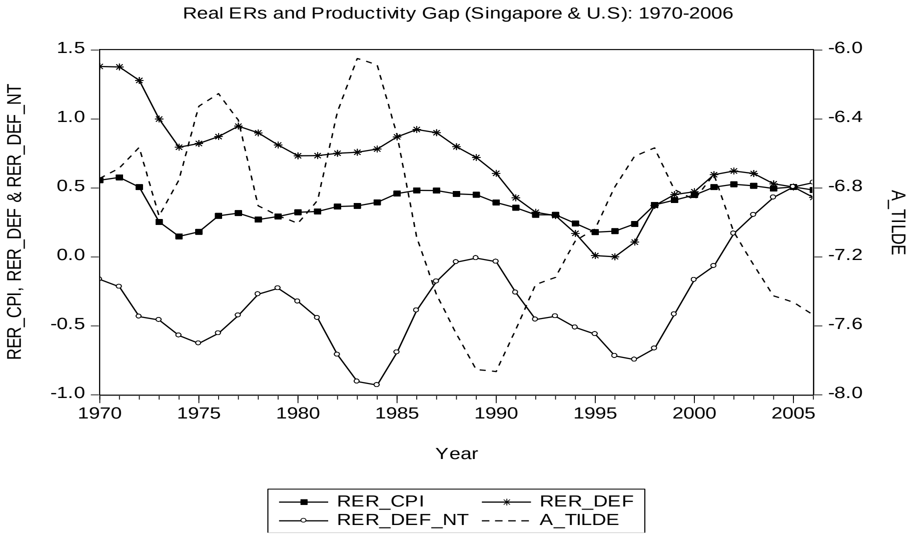
Malaysia



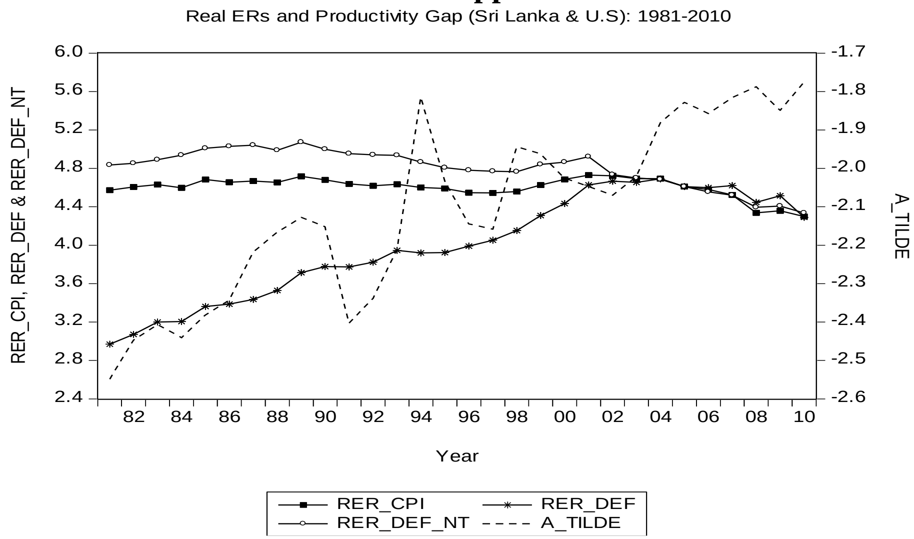
Pakistan



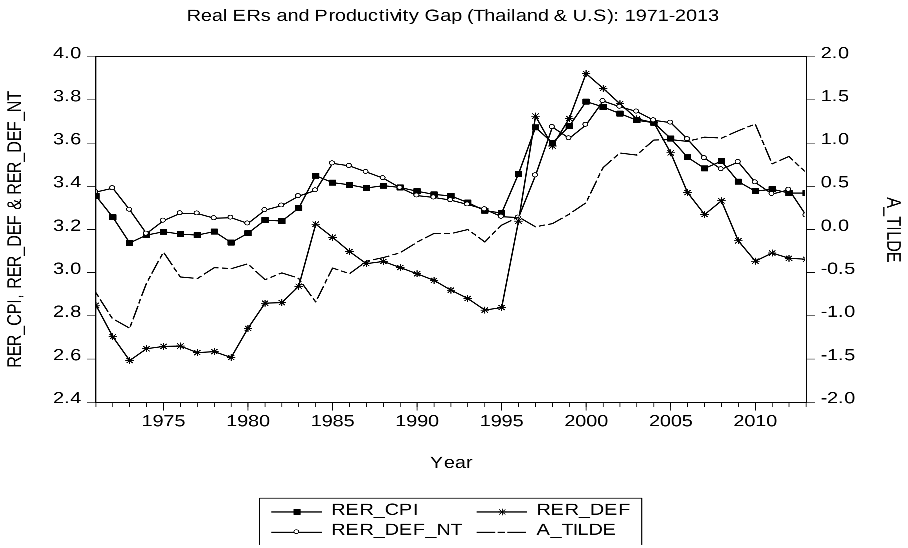
Philippines



Singapore



Sri Lanka


Real ERs and Productivity Gap (Thailand & U.S): 1971-2013
RER_CPI, RER_DEF & RER_DEF_NT
A_TILDE
4.0
3.8
3.6
3.4
3.2
3.0
2.8
2.6
2.4
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
-1.5
-2.0
1975
1980
1985
1990
1995
2000
2005
2010
Year
RER_CPI
RER_DEF
RER_DEF_NT
A_TILDE

Thailand

I start with the single equation cointegration model results. For all three versions of the BS model, the EG test suggests the absence of long-run co-movement between model variables. With regard to determining mean-reversion in errors (obtained through regressing rer on $\tilde{a}$ linearly), the residuals tend to be a unit root process, an undesired behaviour for establishing a valid long-run association amongst model time-series. The null hypothesis of no cointegration between model variables is not rejected using the MacKinnon (1996) critical values.

In contrast, the Error Correction Model (ECM) results suggest otherwise. Against all three versions of the model, the ECM yields a negative and statistically significant EC coefficient. Such findings have strong implications for cointegration between rer and $\tilde{a}$, i.e., rer is making significant adjustments against $\tilde{a}$ shocks to correct short-lived fluctuations, and thus converges to a long-run equilibrium. These findings allow me to proceed next to FMOLS and DOLS regression estimations. The two estimators reveal that the rer_cpi and rer_def_nt based versions of the BS model do not produce results consistent with a BS effect existing for the country. Only the rer_def based model documents valid statistical support in favor of BS hypothesis, as both FMOLS and DOLS estimators produce positive and statistically significant long-run BS coefficients. Thus, two of the three residual-based single equation cointegration tests find a lack of evidence in support of the BS effect for Indonesia.

The multivariate cointegration model under Johansen ML cointegration test specifications concludes that the rank of the cointegrating matrix is either zero or two for all three variants of rer and $\tilde{a}$ relationship, when the model is tested using specification 3 of the test. This implies that no cointegrating is possible between model variables.

TABLE 6.1: Cointegration Tests Results for Indonesia (1976-2013)[42]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[43]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	3	-0.14 [-3.09]	0.41 [1.50]	0.37 [1.07]	No
rer_def	No	3	-0.11[44] [-5.14]	1.62 [2.12]	1.84 [1.86]	Yes
rer_def_nt	No	3	-0.13 [-2.35]	0.14 [0.48]	0.10 [0.27]	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	2	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	1	1	See below
rer_def	1	1	See below
rer_def_nt	1	1	See below

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_cpi	2	Yes	-0.16 [-1.71]	-0.35 [-4.15]	-1.38 [-8.65]	No
rer_def	2	Yes	-0.09 [-0.97]	-0.16 [-4.29]	-3.49 [-9.03]	No
rer_def_nt	2	Yes	-0.07 [-1.26]	-0.21 [-4.61]	-2.31 [-9.09]	No

[42] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[43] t-values are given in squared-brackets.
[44] The test regression also contains first and second lagged-difference of $\tilde{a}$.

However, specification 4 of the test supports the valid existence of cointegration for all three versions of the model. Both Trace and Maximum Eigenvalues indicate a rank of 1, identifying a credible long-run co-movement between rer and $\tilde{a}$. Therefore, I run the VEC model to check the other mandatory conditions for testing for the BS hypothesis.

For neither of the three versions of the model do the VEC results confirm the existence of a valid BS effect. This is because the three conditions necessary to establish an authentic BS effect are not met: (a) the EC_{rer} coefficient is statistically insignificant (except rer_cpi based model which yields marginally significant coefficient), meaning that rer is not making significant adjustments to correct its short-term fluctuations; (b) the condition of weak exogeneity is violated for all three cases of rer, since the $EC_{\tilde{a}}$ coefficient is statistically highly significant; and (c) all the three versions of the model produce a negative BS coefficient, which is highly significant. This implies that $\tilde{a}$ movement of the country against U.S. are pushing up the rer, i.e., the real exchange rate is depreciating against U.S. This behaviour is in conflict with the proposition of the BS effect. Thus, I conclude that the BS hypothesis does not hold true for Indonesia, when the country is tested against U.S.

6.2.2 Japan

The time plots for three alternative measures of Japanese rer and $\tilde{a}$ against the U.S. are given in FIGURE 6.1. From the visual inspection, there are no significant traces of the BS effect in the country. This is because there are less frequent instances of intersection between rer and $\tilde{a}$ throughout the sample period. Furthermore, the two variables are clearly trending in opposite direction; another evidence against the existence of long-run productivity-exchange rate cointegration.

TABLE 6.2: Cointegration Tests Results for Japan (1970-2013)[45]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[46]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	Yes	1	-0.26 [-3.81]	-0.44 [-2.56]	-0.31 [-1.72]	No
rer_def	No	1	-0.23 [-3.37]	-1.64 [-4.56]	-1.54 [-3.73]	No
rer_def_nt	No	1	-0.24 [-3.37]	-0.22 [-1.23]	-0.09 [-0.47]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	2	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

[45] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[46] t-values are given in squared-brackets.

The second panel of TABLE 6.2 reports single equation cointegration test results for the country. Looking at the EG test results, only the rer_cpi based model rejects the null of no cointegration at below the ten percent significance level. Next, I estimate the ECM to test for a long-run relationship between rer and $\tilde{a}$ for all three model variants. The test results yield statistically significant EC coefficients of value -0.26, -0.23, and -0.27 for the rer_cpi, rer_def and rer_def_nt based versions of BS model respectively. This implies that in each period (year), the short-run rer misalignments are significantly adjusted by movements of the respective series, so that the series converge to its long-run equilibrium. Finally, I estimate the long-run BS coefficient. Unfortunately, for all three versions of the model, the coefficient turns out to be negative and/or statistically insignificant. This is evidence against the BS effect for Japan when using the single equation cointegration approach.

Now, I move towards testing for a possible cointegrating relationship between rer and $\tilde{a}$ using the Johansen ML multivariate cointegration test procedure. For all three variants of the model, a rank of zero or two is obtained against both of the test specifications. This proves that the country's productivity differential does not have a good tendency for explaining the long-run dynamics of rer against the U.S. Thus, like the single equation results, the vector error correction model also does not find evidence for the BS effect for Japan.

6.2.3 Korea

From FIGURE 6.1, it is clear that the three rer and $\tilde{a}$ series for Korea and the U.S. display quite dissimilar trend movements over the sample period. This is counter to what one would expect of a long-run association predicted under the BS hypothesis. Also, the two series display quite infrequent intersections, a pattern rejecting the likely presence of a long-term cointegrating relationship between rer and $\tilde{a}$.

TABLE 6.3: Cointegration Tests Results for Korea (1970-2013)[47]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(0)**	Inconclusive
rer_def	Yes	I(1)***	I(0)**	Inconclusive
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[48]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	Yes	1	-0.39 [-3.35]	-0.04 [-1.62]	-0.05 [-1.66]	No
rer_def	Yes	4	-0.52 [-2.78]	-0.20 [-3.77]	-0.21 [-2.94]	No
rer_def_nt	No	4	-0.43 [-2.54]	0.02 [0.49]	0.01 [0.25]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	2	0	No
rer_def	0	0	No
rer_def_nt	1	1	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
rer_def_nt	3	Yes	-0.42 [-3.48]	0.10 [0.43]	0.13 [1.28]	No

[47] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[48] t-values are given in squared-brackets.

For the cases of the CPI and GDP deflator based models, but not the GDP deflator (nontradables) based model, the single equation test (under EG specifications) rejects the null hypothesis of no cointegration at the 10 percent significance level. The associated tau-values, when compared to the MacKinnon (1996) critical values, suggest long-run causality between $\tilde{a}$ and rer for Korea against the U.S. This is confirmed when I try to verify the long-run association between rer and $\tilde{a}$ through the ECM, as the test supports cointegration for all three model variants. The EC coefficient takes values of -0.39, -0.52, and -0.43 for CPI deflated, GDP deflator and GDP deflator (nontradables) based models respectively, with high statistical confidence. But the long-run BS effect is estimated to be negative and/or statistically insignificant for all three models using the FMOLS and DOLS estimators. Thus, Korea's inter-country productivity gap with the U.S. is associated with depreciation in the rer, instead of appreciation, a result that contradicts the BS hypothesis.

Testing the BS model in the context of a vector error correction model, I obtain one valid cointegrating vector for the case of nontradables prices based model, under specification 3 of the model. However, the corresponding VEC model does not produce a statistically significant BS coefficient. Thus the BS hypothesis does not find support in the case of Korea.

6.2.4 Malaysia

Starting with the visual assessment in FIGURE 6.1, the trend movements of $\tilde{a}$ and three rer series are very illuminating. This is primarily because of the $\tilde{a}$ behavior, which sharply falls for the first decade of the sample period and then suddenly rises during the last decade. For almost the entire sample period, the three indices of rer and $\tilde{a}$ series are not co-moving in a common direction. Furthermore, there are only a moderate number of instances where the two types of series intersect over the data period.

TABLE 6.4: Cointegration Tests Results for Malaysia (1980-2013)[49]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)**	I(1)***	I(1)
$\tilde{a}$	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)

Single Equation Cointegration Approach[50]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC[51] Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	4	-0.06 [-1.28]	-	-	No
rer_def	No	2	-0.15 [-2.08]	-0.30 [-2.33]	-0.45 [-1.63]	No
rer_def_nt	No	2	-0.22 [-2.57]	-0.20 [-2.84]	-0.19 [-1.33]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	1	1	See below
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
rer_cpi	3	Yes	-0.67 [-4.08]	-0.71 [-1.24]	-0.17 [-5.39]	No

[49] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[50] t-values are given in squared-brackets.
[51] The three test regressions also contain first and second lagged-difference of $\tilde{a}$.

Malaysia's cointegration test results are provided in TABLE 6.4. Starting with the single equation cointegration model (under the EG test specifications), there are no signs of long-run cointegration for any of the measures of rer and their long-run $\tilde{a}$ determinants. The ECM results also do not produce evidence of co-movement between rer and $\tilde{a}$ for *rer_cpi*. In contrast, the GDP deflator and GDP deflator (nontradables) based version of the models show significant traces of a valid cointegration relationship among the model variables. However, when the models are tested for a long-run BS effect, the FMOLS and DOLS estimators yield negative coefficients, invalidating the BS hypothesis for the country.

Similar results are produced by the vector error correction model. For the rer and $\tilde{a}$ variables, neither the Trace or Max Eigen statistics indicate cointegration except for the case of the rer_cpi based model for specification 4, where the test indicates a rank of one. I follow this up by estimating the VEC model. However, the VEC results do not support the BS hypothesis, as the estimate for the long-run BS coefficient is negative and statistically significant. Overall, these findings are consistent with the results yielded through the single equation cointegration model. Hence, like previously discussed countries, the BS hypothesis is not supported in the case of Malaysia.

6.2.5 Pakistan

In FIGURE 6.1, the time plots of the three measures of rer and $\tilde{a}$ are given for Pakistan against the U.S. From the visual assessment, it seems like there is little possibility of a long-run association between the two series. After the year 1987, for almost the entire sample period, the two series are trending in a dissimilar direction. Furthermore, their irregular intersection during this period points to their inability to establish long-run cointegration.

TABLE 6.5: Cointegration Tests Results for Pakistan (1973-2008)[52]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	I(1)**	I(1)**	I(1)
rer_def_nt	Yes	I(1)**	I(1)*	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[53]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	1	-0.03 [-0.91]	-	-	No
rer_def	No	1	-0.02 [-0.70]	-	-	No
rer_def_nt	No	1	-0.05[54] [-1.18]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

52 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
53 t-values are given in squared-brackets.
54 The three test regression also containS first and second lagged-difference of $\tilde{a}$.

The single equation cointegration test results for Pakistan are given in TABLE 6.5. The EG test and ECM produce common results. The tau-statistics generated by the EG test are unable to reject the null hypothesis of no cointegration for all three versions of the BS model. The same is true for the ECM, where the test produces a statistically insignificant EC coefficient. Thus, the BS effect does not appear to hold for Pakistan, at least according to the single equation cointegration test findings.

Next I move towards investigating a possible cointegrating relationship between the three rer series with $\tilde{a}$ using the Johansen ML cointegration test procedure. In line with my single equation cointegration test results, both cases of the Johansen model (3 and 4) do not support the valid existence of a causal effect between model variables for Pakistan. For all three variants of the model, a rank of zero or two is obtained. This indicates that the country's productivity differential does not have good explanatory power for explaining the long-run dynamics of its real exchange rate against the U.S. Thus, similar to the single equation test findings, I find no evidence in favor of the BS effect for Pakistan.

6.2.6 Philippines

The rer and $\tilde{a}$ time plots for the Philippines are suggestive of a BS effect for the country. The $\tilde{a}$ series is co-moving with the three measures of rer and is closely co-moving with the GDP deflator based measure. Furthermore, the rer series is found to be trending with an efficient speed of adjustment which may cause convergence towards its long-run equilibrium.

From the EG single equation test results, statistical support in favor of a cointegrating relationship can be found for the CPI and GDP deflator (nontradables) based models, but not for the *rer_def* based model.

TABLE 6.6: Cointegration Tests Results for Philippines (1971-2013)[55]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(0)***	I(1)***	Inconclusive
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[56]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of *rer*	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	Yes	4	-0.30 [-2.95]	0.06 [0.54]	0.20 [1.14]	No
rer_def	No	4	-0.06 [-1.52]	-	-	No
rer_def_nt	Yes	3	-0.31 [-3.35]	-0.01 [-0.12]	0.10 [0.61]	No

Multivariate Cointegration Approach

Version of *rer*	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	1	1	See below
rer_def	1	0	See below
rer_def_nt	1	0	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Version of *rer*	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_cpi	3	Yes	-0.29 [-3.00]	0.14 [0.78]	0.06 [0.44]	No
rer_def	3	Yes	-0.07 [-1.97]	-0.03 [-0.91]	-0.10 [-0.08]	No
rer_def_nt	2	Yes	-0.31 [-3.49]	-0.04 [-0.21]	-0.07 [-0.58]	No

55 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
56 t-values are given in squared-brackets.

For these two models, the tau statistics generated by the EG test exceed the Mackinnon critical values at the 10 percent significance level. On estimating the error correction dynamics of the three variants of the BS model, the rer series turn out to be adjusting significantly to correct short-run errors/deviations, returning the series to its long-run equilibrium, except for the case of GDP deflator based rer. However, on estimating the long-run BS coefficient using the FMOLS and DOLS estimators, a positive and statistically significant coefficient is found only for the GDP deflator based model. Evidence of a BS effect is not found using the other two rer series. Thus, the single equation cointegration test results do not indicate a robust relationship between rer and $\tilde{a}$ for the Philippines.

The multivariate cointegration model finds moderate support of a cointegrating relationship between the respective series. Either the Trace statistic or Max Eigenvalue statistic or both produces a rank of one for the three model variants under specification 3 of the Johansen ML test. On obtaining support in favor of a significant long-run co-movement between rer and $\tilde{a}$, I estimate the VEC model. But the VEC estimates rule out the possibility of a BS effect for the country. All three models successfully meet the first two conditions for a valid BS effect, associated with error correction adjustments of rer and weak exogeneity of $\tilde{a}$, but they do not produce statistically significant long-run BS coefficients. Hence, the BS model does not hold valid for the Philippines.

6.2.7 Singapore

The time plot for the Singaporean rer and $\tilde{a}$ series are given in FIGURE 6.1. The graphs are quite illuminating as the trend patterns of two of the three rer series (the GDP deflator-nontradables based version excepted) and $\tilde{a}$ are seemingly compatible with the theoretical predictions of the BS model. The $\tilde{a}$ series is regularly intersecting with the CPI-deflated and GDP

TABLE 6.7: Cointegration Tests Results for Singapore (1970-2006)[57]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(0)***	I(0)***	I(0)
rer_def	Yes	Greater than I(1)	I(0)**	Inconclusive
rer_def_nt	Yes	I(0)**	I(0)**	I(0)
$\tilde{a}$	Yes	I(1)**	I(0)**	Inconclusive

Single Equation Cointegration Approach[58]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	Yes	2	-0.24 [-1.74]	-0.06 [-1.13]	-0.09 [-1.24]	No
rer_def	Yes	2	-0.13 [-2.51]	0.30 [1.76]	0.43 [2.01]	Yes
rer_def_nt	No	2	0.08 [1.50]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	2	0	No
rer_def	2	2	No
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	1	0	See below
rer_def	1	0	See below
rer_def_nt	0	0	No

Vector Error Correction Model

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
rer_cpi	1	Yes	-0.33 [-4.77]	-0.27 [-0.88]	0.10 [1.61]	No
rer_def	1	Yes	-0.15 [-3.99]	-0.03 [-0.28]	0.74 [2.98]	Yes

[57] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[58] t-values are given in squared-brackets.

deflator based rer series during the sample period. But the two series have large swings (particularly the $\tilde{\alpha}$ series). In contrast, productivity movements do not appear to be systematically related with nontradables prices based rer movements for almost the entire sample period. Such mixed data characteristics leaves me uncertain regarding the long-run association between $\tilde{\alpha}$ and rer.

On conducting single equation cointegration tests, the model finds weak empirical evidence in support of cointegration. On conducting the EG cointegration test, both CPI and GDP deflator based models reject the null hypothesis of no cointegration with appropriate statistical significance. Such is the not the case for the GDP deflator (nontradables) model. Nor does the ECM find support for cointegration for the latter version of the *rer*. When I estimate the ECM and the long-run BS coefficient, the GDP deflator based model produces a statistically significant and positive BS coefficient. In contrast, the FMOLS and DOLS estimates of the BS coefficient are insignificant for the CPI based model. Thus, the single equation cointegration model produces evidence for the BS effect for only one of the three *rer* variants.

Similar statistical support for the BS effect is obtained from the multivariate cointegration model. The Trace statistic produces a rank of one for the CPI and GDP deflator based models under specification 4 of the test. However, the necessary conditions in support of the BS hypothesis are found only for the GDP deflator based model. A positive and statistically significant BS coefficient of 0.73 is estimated for this model. Insufficient evidence is found for the other two versions of *rer*. As only one of the three *rer* versions in both the single and multiple equation models meet the appropriate conditions, I conclude that there is only weak support for the BS hypothesis for Singapore.

6.2.8 Sri Lanka

From a visual inspection of Sri Lanka's time plots of the three variants of rer and $\tilde{a}$ against the U.S., it appears that these series provide support for the BS effect at the outset. This is because the two series are trending in a similar direction throughout the sample period, a feature making the plausible long-run association likely to hold. Moreover, frequent intersection of $\tilde{a}$ and the GDP deflator based measures of rer provides strong evidence for their long-run association.

The country results are reported in TABLE 6.8. In contrast to the first impression from the time series plots, the results from the single equation cointegration tests do not provide support for cointegration for two of the three *rer* models. The exception is the CPI based version of the model, where the estimated ECM displays a negative and statistically significant EC coefficient. However, the BS hypothesis is not supported by the corresponding FMOLS and DOLS estimates. Both coefficients are negative, and one is significantly negative. Taken together, the results from the single equation analysis provides consistent evidence against the BS hypothesis for Sri Lanka.

The results from the multivariate cointegration model lead to a slightly different conclusion regarding the existence of a cointegrating relationship among the variables. Under specification 3 of the test, the Trace and Max Eigenvalue statistics indicate that the variables are not cointegrated. However, under specification 4 of the test, both tests provide evidence of a cointegrating relationship for all three model variants. Nevertheless, the estimated VEC model is discouraging with respect to the BS hypothesis. The EC_{rer} coefficients are all insignificant, while the $EC_{\tilde{a}}$ coefficients are significant. And the BS coefficients are wrong-signed and significant in all three models. This indicates that the country's productivity differences with the U.S. are associated with a decrease in relative prices (nontradables), instead of an increase, exactly the opposite of what the BS hypothesis predicts.

TABLE 6.8: Cointegration Tests Results for Sri Lanka (1981-2010)[59]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	I(1)***	I(1)
rer_def	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	I(0)***	I(0)

Single Equation Cointegration Approach[60]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	2	-0.33 [-3.73]	-0.26 [-1.97]	-0.21 [-1.43]	No
rer_def	No	1	-0.05 [-1.38]	-	-	No
rer_def_nt	No	1	-0.08 [-0.82]	-	-	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	0	0	No
rer_def	2	0	No
rer_def_nt	2	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	1	1	See below
rer_def	1	1	See below
rer_def_nt	1	1	See below

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_cpi	1	Yes	-0.00 [-0.41]	-0.17 [-5.01]	-6.00 [-5.07]	No
rer_def	0	Yes	0.02 [0.75]	-0.15 [-5.05]	-6.83 [-5.14]	No
rer_def_nt	0	Yes	-0.01 [-1.11]	-0.07 [-3.70]	-8.98 [-3.74]	No

[59] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[60] t-values are given in squared-brackets.

6.2.9 Thailand

The time plots for the three alternative measures of Thailand's rer and $\tilde{a}$ against the U.S. are given in FIGURE 6.1. From visual inspection, there are significant traces of the BS effect in the country until year 1996, as the two types of series closely trend together. However, after this period, all three variants of rer and $\tilde{a}$ move in opposite directions. Also, there are less frequent instances of intersection between rer and $\tilde{a}$ through this period.

TABLE 6.9 reports cointegration test results for the country. From EG test results, the tau-statistics tend to reject the null hypothesis of no cointegration at better than 5 percent statistical significance for the cases of rer_cpi and rer_def (only). However, the ECM results for all three models favors valid convergence of rer towards its long-run equilibrium. The test results yield negative and statistically significant EC coefficients, revealing significant adjustments on the part of the respective rer variants to correct deviations from long-run equilibrium. Nevertheless, the final stage results are rather disappointing. The two cointegration regression estimators (FMOLS and DOLS) produce a negative long-run BS coefficient for all model variants. Thus, there is again no evidence of a BS effect for Thailand based on an analysis of the single equation cointegration models.

Now, I move towards exploring the possible cointegrating relationship between the three rer series with $\tilde{a}$ using the Johansen ML Cointegration test procedure. For all three versions of the model under specification 3 of the test, either the Trace statistic or the Max Eigenvalue statistic or both indicate the existence of one cointegrating vector. No evidence of a cointegrating relationship is found under specification 4 of the model.

TABLE 6.9: Cointegration Tests Results for Thailand (1971-2013)[61]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	I(1)***	Greater than I(1)	Inconclusive
rer_def	Yes	I(1)***	I(1)***	I(1)
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[62]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	Yes	1	-0.17 [-1.96]	-0.28 [-4.76]	-0.33 [-5.41]	No
rer_def	Yes	1	-0.14[63] [-1.76]	-0.57 [-4.52]	-0.67 [-4.66]	No
rer_def_nt	No	1	-0.17 [-1.91]	-0.49 [-6.78]	-0.55 [-7.13]	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	1	1	See below
rer_def	1	0	See below
rer_def_nt	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	0	0	No
rer_def	0	0	No
rer_def_nt	0	0	No

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_cpi	0	Yes	-0.09 [-1.17]	-0.54 [-2.23]	-0.37 [-3.91]	No
rer_def	0	Yes	-0.06 [-0.81]	-0.28 [-2.48]	-0.79 [-3.86]	No
rer_def_nt	0	Yes	-0.07 [-1.05]	-0.54 [-2.80]	-0.58 [-5.71]	No

61 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
62 t-values are given in squared-brackets.
63 The test regression also contains first lagged-difference of $\tilde{a}$.

Returning to specification 3, estimation of the VEC model does not provide support for the BS hypothesis for any of the model variants, and for none of the necessary conditions. The error correction coefficients in the *rer* equations are not significant, while the error correction coefficients in the productivity equation are significant. Further, the BS coefficients are all estimated to be negative and statistically significant. Thus, similar to the single equation test findings, here also I reject the possible existence of a BS effect for the country.

6.2.10 Summary of Individual Country Studies

TABLE 6.10 summarizes the preceding results for the individual countries in my sample. The table documents both single equation and multivariate cointegration test results. With a few minor exceptions, the results consistently indicate rejection of the BS hypothesis. Indonesia, the Philippines, and Singapore are the only countries for which there is even slight evidence for the BS hypothesis. Interestingly, in these few instances, it always occurs using the GDP deflator based version of the model. This is in line with the existing BS literature on Asia which has used a gross output deflator (GDP or GNP) based real exchange rate as the model regressand (Bahmani-Oskooee and Rhee, 1996; Chinn, 2000; Bahmani-Oskooee and Nasir, 2004; Thomas and King, 2008). For the other countries, and the other model variants, there is absolutely no evidence in support of the BS hypothesis. To conclude, when I use finer sectoral divisions in estimating the BS model, it does not cause me to modify my previous conclusions about the lack of support for the BS hypothesis.

TABLE 6.10: Does Balassa-Samuelson Effect Hold? Summary of Results by Country

Country	Individual Results				Country Summary
	Version of rer	*Single Equation Cointegration Method*	*Multivariate Cointegration Method*		
			Case 3	*Case 4*	
Indonesia (1976-2013)	rer_cpi	No	No	No	Mixed
	rer_def	Yes	No	No	
	rer_def_nt	No	No	No	
Japan (1970-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Korea (1970-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Malaysia (1980-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Pakistan (1973-2008)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Philippines (1971-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Singapore (1970-2006)	rer_cpi	No	No	No	Mixed
	rer_def	Yes	No	Yes	
	rer_def_nt	No	No	No	

Country	Individual Results				Country Summary
	Version of rer	*Single Equation Cointegration Method*	*Multivariate Cointegration Method*		
			Case 3	*Case 4*	
Sri Lanka (1981-2010)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	
Thailand (1971-2013)	rer_cpi	No	No	No	No
	rer_def	No	No	No	
	rer_def_nt	No	No	No	

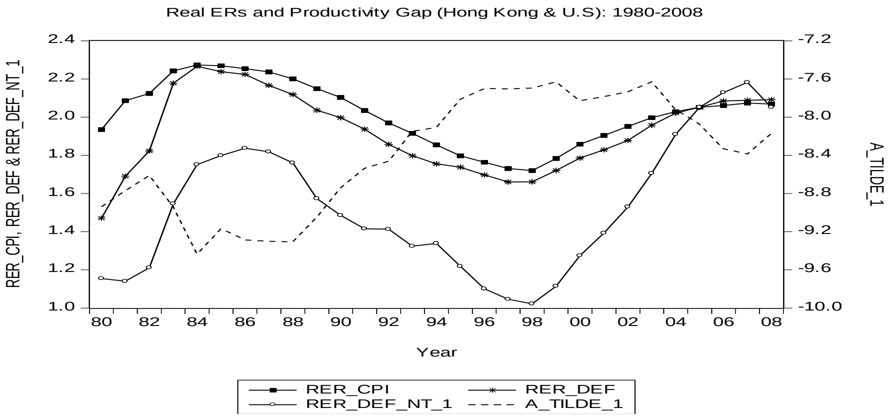
6.2.11 Hong Kong

As I elaborated in Chapter Five, Hong Kong needs to be handled in a different manner than the other countries because Hong Kong does not follow a unique sectoral division. Accordingly, my empirical analysis for the country applies four different approaches to mapping Hong Kong's industrial classifications to tradable and nontradable sectors.

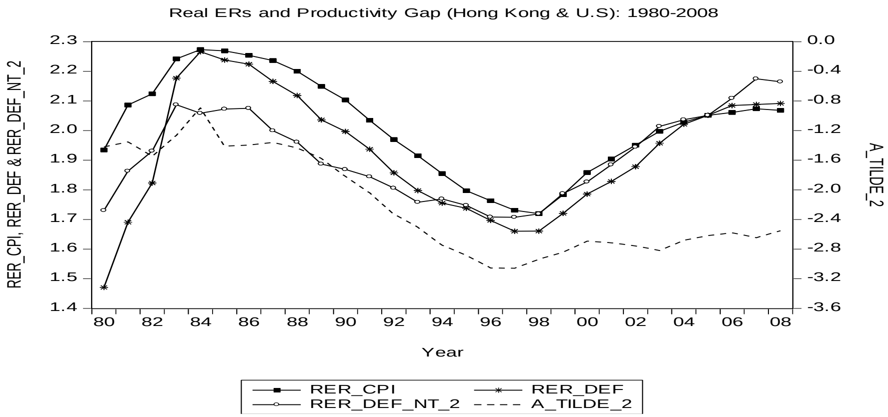
Starting with the rer and $\tilde{a}$ relationship under the first type of sectoral division (Case1: Hong Kong_1), there is no significant long-run association between rer and $\tilde{a}$. The tau statistics associated with the cointegration test could not reject the null hypothesis of no cointegration with acceptable statistical significance. The same is true for my second single equation cointegration test (ECM), where all three model variants yield statistically insignificant EC coefficients. Thus, there is no evidence for the BS effect for Hong Kong, at least according to the single equation cointegration models.

The results generated by the multivariate cointegration tests do not differ substantially from the single equation tests. The Johansen ML tests find no evidence of a cointegrating vector for the CPI and GDP deflator based models. The results are slightly different for the GDP deflator (nontradables) model. Under specification 4, both Trace and Max Eigenvalue statistics suggest one valid cointegrating vector when the *rer_def_nt_1* measure of real exchange rates is used. However, the corresponding VEC estimates are not supportive. None of the three conditions necessary to establish a reliable BS effect are met. The EC_{rer} and $EC_{\tilde{a}}$ coefficients have reversed properties from what would be expected were the BS hypothesis valid. The former is statistically insignificant, while the latter is significant and negative. In addition, the long-run BS coefficient is wrong-signed and statistically significant. Thus, Case 1 for Hong Kong does not provide support for the BS hypothesis.

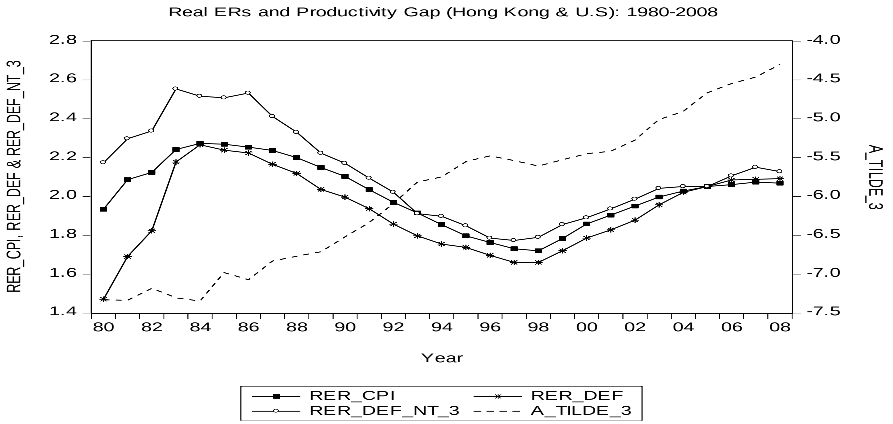
FIGURE 6.2: Plots for Real Exchange Rates and Sectoral Productivity Differentials against U.S. (Hong Kong: 1980-2008)



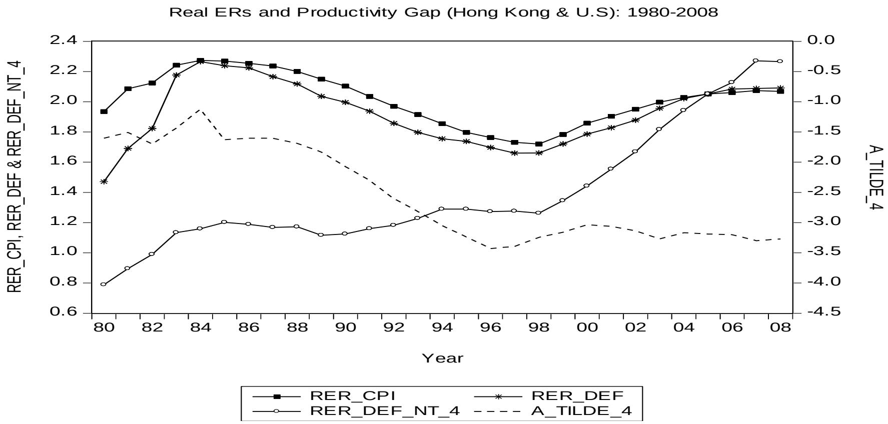
Hong Kong_1[64]



Hong Kong_2[65]



Hong Kong_3[66]



Hong Kong_4[67]

[64] Construction is the only nontradable sector here. Rest of all the sectors are treated as tradables.
[65] Mining, utilities, construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.
[66] Mining, utilities and construction are nontradable sectors whereas rest of all the sectors are treated as tradables.
[67] Construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.

TABLE 6.11: Cointegration Tests Results for Hong Kong: Case 1 (1980-2008)[68]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	Greater than I(1)	I(0)**	Inconclusive
rer_def	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_nt\_1$	Yes	Greater than I(1)	I(1)**	Inconclusive
$\tilde{a}\_1$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[69]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	3	-0.11 [-0.75]	-	-	No
rer_def	No	3	-0.02 [-0.20]	-	-	No
$rer\_def\_nt\_1$	No	3	-0.11 [-1.63]	-	-	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	2	2	No
rer_def	2	2	No
$rer\_def\_nt\_1$	0	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	2	2	No
rer_def	2	2	No
$rer\_def\_nt\_1$	1	1	See below

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}\_1}$	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$rer\_def\_nt\_1$	2	Yes	-0.14 [-0.63]	-1.09 [-3.03]	-0.74 [-14.65]	No

[68] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[69] t-values are given in squared-brackets.

Next, I discuss the empirical results for other three classifications of Hong Kong together, as they are quite alike. The rer and $\tilde{a}$ time plots for second, third and fourth classifications of Hong Kong's sectoral divisions are given in FIGURE 6.2. There appears to be some support for the possibility of a cointegrating relationship between the real exchange rate and productivity variables, depending on the particular sectoral classifications one uses.

None of the three *rer* GDP deflator (nontradables) variants appear to be cointegrated with the productivity variable according to the single equation, EG test. The ECM tests for cointegration are mixed, with some of the EC coefficients indicating cointegration, and some not. In particular, the ECM models for the variables *rer_cpi* (Case 2), *rer_def* (Case 2), *rer_cpi* (Case 3), *rer_def* (Case 3), *rer_def_nt_3* (Case 3), *rer_cpi* (Case 4) and *rer_def* (Case 4) all find evidence of a cointegrating relationship. However, when FMOLS and DOLS are used to estimate the long-run relationship between the respective real exchange rate and productivity variables, evidence for the BS hypothesis is only found in the first case, *rer_cpi* (Case 2) and *rer_def* (Case 2) and the fourth case, *rer_cpi* (Case 4) and *rer_def* (Case 4), with mixed support for *rer_def_nt_2* (Case 2). The BS coefficients in all the other cases are negative, with most of them being significant.

The results from the multivariate cointegration models produce similar results. There is some evidence for all three cases of a cointegrating relationship. All three sectoral classifications have at least one instance where one cointegrating vector is identified. However, in none of the cases does one find support for the BS hypothesis. The long-run BS coefficient is statistically insignificant or negative or both for all cases, except rer_def, Case 2, yielding positive and statistically significant BS coefficient. Also, the condition of weak exogeneity is frequently violated as the error correction coefficient ($EC_{\tilde{a}}$) is statistically significant in Cases 2 and 4. I conclude that a valid BS effect cannot be found for any of the four sectoral divisions for Hong Kong.

TABLE 6.12: Cointegration Tests Results for Hong Kong: Case 2 (1980-2008)[70]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	Greater than I(1)	I(0)**	Inconclusive
rer_def	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_nt\_2$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_2$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[71]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	4	-0.17[72]	0.22	0.25	Yes
			[-3.38]	[5.23]	[4.74]	
rer_def	No	4	-0.12[73]	0.16	0.27	Yes
			[-1.98]	[2.33]	[4.05]	
$rer\_def\_nt\_2$	No	4	-0.14	0.08	0.13	Mixed
			[-1.94]	[1.33]	[1.98]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	1	1	See below
rer_def	1	1	See below
$rer\_def\_nt\_2$	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	2	2	No
rer_def	2	2	No
$rer\_def\_nt\_2$	2	2	No

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}\_2}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_cpi	3	Yes	-0.11	-0.32	-0.07	No
			[-5.08]	[-2.20]	[-1.33]	
rer_def	3	Yes	-0.00	-0.03	-4.00	No
			[-4.90]	[-3.54]	[-5.40]	
$rer\_def\_nt\_2$	3	Yes	-0.01	-0.02	-3.42	No
			[-4.31]	[-1.97]	[-5.46]	

[70] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[71] t-values are given in squared-brackets.
[72] The test regression also contains first lagged-difference of $\tilde{a}$.
[73] The test regression also contains first and second lagged-differences of $\tilde{a}$.

TABLE 6.13: Cointegration Tests Results for Hong Kong: Case 3 (1980-2008)[74]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	Greater than I(1)	I(0)**	Inconclusive
rer_def	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_nt\_3$	Yes	I(1)**	I(1)***	I(1)
$\tilde{a}\_3$	Yes	I(0)**	I(0)***	I(0)

Single Equation Cointegration Approach[75]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
rer_cpi	No	2	-0.15[76] [-3.26]	-0.10 [-1.89]	-0.15 [-2.59]	No
rer_def	No	4	-0.09[77] [-1.89]	-0.05 [-0.76]	-0.13 [-1.82]	No
$rer\_def\_nt\_3$	No	4	-0.17 [-2.75]	-0.17 [-2.60]	-0.23 [-3.07]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	1	1	See below
rer_def	1	1	See below
$rer\_def\_nt\_3$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_cpi	2	0	No
rer_def	1	1	See below
$rer\_def\_nt\_3$	2	2	No

[74] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[75] t-values are given in squared-brackets.
[76] The test regression also contains first lagged-difference of $\tilde{a}$.
[77] The test regression also contains first and second lagged-differences of $\tilde{a}$.

Vector Error Correction Model[78]

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}\_3}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_cpi	1	Yes	-0.08 [-2.10]	0.19 [1.22]	-0.03 [-0.43]	No
$rer\_def\_nt\_3$	3	Yes	-0.09 [-5.84]	0.07 [1.32]	0.28 [3.01]	Yes
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_def	3	Yes	-0.17 [-2.57]	0.18 [1.33]	-0.11 [-0.24]	No

[78] t-values are given in squared-brackets.

TABLE 6.14: Cointegration Tests Results for Hong Kong: Case 3 (1980-2008)[79]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_cpi	Yes	Greater than I(1)	I(0)**	Inconclusive
rer_def	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_nt\_4$	Yes	Greater than I(1)	I(1)**	Inconclusive
$\tilde{a}\_4$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[80]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of rer	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
rer_cpi	No	4	-0.13[81] [-2.61]	0.17 [3.76]	0.20 [3.61]	Yes
rer_def	No	4	-0.11[82] [-1.76]	0.12 [1.73]	0.20 [2.93]	Yes
$rer\_def\_nt\_4$	No	4	-0.01 [-0.28]	-	-	No

Multivariate Cointegration Approach

Version of rer	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_cpi	1	1	See below
rer_def	1	1	See below
$rer\_def\_nt\_4$	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_cpi	2	2	No
rer_def	2	2	No
$rer\_def\_nt\_4$	2	2	No

Vector Error Correction Model

Version of rer	Lags	White Noise Residuals	EC_{rer}	$EC_{\tilde{a}\_4}$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_cpi	3	Yes	-0.09 [-5.81]	-0.26 [-2.07]	-0.12 [-2.45]	No
rer_def	3	Yes	-0.04 [-5.28]	-0.09 [-3.30]	-0.86 [-5.21]	No
$rer\_def\_nt\_4$	3	Yes	-0.04 [-2.04]	-0.11 [-2.52]	-0.98 [-4.66]	No

79 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
80 t-values are given in squared-brackets.
81 The test regressions also contain first and second lagged-differences of $\tilde{a}$.
82 The test regressions also contain first and second lagged-differences of $\tilde{a}$.

6.3 Panel Data Estimation Results

I start my panel data estimations by formally testing the model variables for their order of integration using panel unit root tests. I am testing the variables using Fisher-ADF and Fisher-PP unit root tests and Hadri stationarity test.

The results are reported in the first panel of TABLE 6.15. For the GDP deflator based measure of real exchange rates, there is unanimous support from all three tests, indicating the rer_def series is a unit root process in levels. The empirical support for the CPI and GDP deflator (nontradables) based rer measures is somewhat weaker, as these variables are stationary in levels according to the Fisher-ADF test. In contrast, the results of the Hadri and Fisher-PP tests provide evidence at the one percent level that these variables are integrated of order one. Overall, I conclude that all three *rer* series to be unit root processes in levels. For the $\tilde{a}$ series, two unit root tests indicate that this series is level stationary. On the other hand, the Hadri stationarity test indicates that the order of integration is greater than one. Though the test results are inconsistent, I conclude that the productivity variable is stationary in levels.

The second and third panels of TABLE 6.15 display the test results for the Pedroni residual based cointegration tests and the Johansen Fisher panel cointegration tests. With regard to the Pedroni cointegration tests, I opted for automatic lag selection. A majority of the Pedroni cointegration test statistics (at least 5 out of 7 tests) fail to reject the null hypothesis of no cointegration between model variables. Similar lack of evidence for cointegration was found using specification 3 for the Johansen Fisher panel cointegration tests. However, specification 4 yielded evidence for cointegration from both Trace and Max Eigenvalue statistics for all three *rer* versions. Accordingly, I estimated panel PFMOLS and panel PDOLS regressions for these variables. All the corresponding BS coefficient estimates were negative, with several being significant.

Thus, in the end, I was unable to find evidence for the BS hypothesis from the pooled data analysis for any of the three *rer* variants of the model.

6.3.1 Summary of Panel Data Results

The panel data estimators yield the same conclusions as those from the individual country studies. In fact, relative to the individual country estimates, where there was at least occasional support, the panel results produced no evidence in favor of the BS hypothesis. Both single equation and multivariate panel estimators came up empty-handed in the search for a BS effect. Thus, I conclude that using better sectoral categorizations to capture tradables and nontradables – as some have advocated (Rother, 2000; Mihaljek et al., 2003, 2004; Coricelli and Jazbec, 2004; Egert, 2003, 2004, 2005; Gibson, 2008) -- does not change the conclusions one obtains using cruder sectoral divisions. These findings call for revisiting the BS theory from the perspective of theoretical specifications of the model. Subsequent chapters of this dissertation pursue this line of inquiry.

TABLE 6.15: Summary of Results for Panel Unit Root and Cointegration Tests for Balassa-Samuelson Effect[83,84]

Panel Unit Root Test Results (Order of Integration as Determined by)

Variables	Fisher-ADF	Fisher-PP	Hadri	Conclusion
rer_cpi	I (0)**	I (1)***	I (1)***	I (1)
rer_def	I (1)***	I (1)***	I (1)***	I (1)
rer_def_nt	I (0)*	I (1)***	I (1)***	I (1)
$\tilde{a}$	I (0)**	I (0)**	Greater than I (1)	I (0)

Pedroni Panel Cointegration Test Results[85]

	Common AR Coefficients (Within Dimension)				*Individual AR Coefficients (Between Dimension)*			
Version of rer	Panel v Statistics	Panel ρ Statistics	Panel PP Statistics	Panel ADF Statistics	Group ρ Statistics	Group PP Statistics	Group ADF Statistics	Does BS Effect Hold?
rer_cpi	0.20	-0.21	-1.13	-2.56***	0.92	-0.19	-3.91***	No
rer_def	1.37*	-0.32	-1.01	-1.58*	1.15	0.38	-0.69	No
rer_def_nt	-0.90	0.50	-0.13	-0.98	1.29	0.46	-1.63*	No

83 ***, ** and * are representing significance of sample statistics at 1%, 5% and 10% levels respectively.
84 Hong Kong is omitted from panel estimations.
85 Pedroni panel cointegration is a test for null of no cointegration in both homogenous and heterogeneous panels. The test statistics are standardized and asymptotically normally distributed. See Pedroni (1995, 1999) for further details.

Johansen-Fisher Panel Cointegration Test Results[86,87]

Version of *rer*	Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigen Stat)	Does BS Effect Hold?
Case 3: Intercept (no trend) in cointegrating equation and VAR			
rer_cpi	2	2	No
rer_def	2	2	No
rer_def_nt	2	2	No
Case 4: Intercept and trend in cointegrating equation-no trend in VAR			
rer_cpi	1	1	See below
rer_def	1	1	See below
rer_def_nt	1	1	See below

Results for Panel FMOLS and DOLS[88] **Estimators**
Long-Run Cointegrating Vectors for Balassa-Samuelson Effect

Version of *rer*	**Estimator**	**BS Coefficient**[89]	**Does BS Effect Hold?**
rer_cpi	PFMOLS	-0.05 [-1.19]	No
	PDOLS	-0.08 [-1.83]	No

[86] The test is Maximum likelihood based rank test.
[87] Lag selection is done through SIC under panel VAR.
[88] Lead = Lag = 1.
[89] t-values are given in squared-brackets.

Results for Panel FMOLS and DOLS[88] Estimators Long-Run Cointegrating Vectors for Balassa-Samuelson Effect			
Version of rer	**Estimator**	**BS Coefficient**	**Does BS Effect Hold?**
rer_def	PFMOLS	-0.05 [-0.39]	No
	PDOLS	-0.09 [-0.68]	No
rer_def_nt	PFMOLS	-0.10 [-2.23]	No
	PDOLS	-0.12 [-2.36]	No

TABLE 6.16: Does Balassa-Samuelson Effect Hold? Summary of Results for Panel Cointegration Tests

Version of rer	Test of Cointegration: Pedroni Residual Based Panel Cointegration Test	Johansen-Fisher Panel Cointegration Test: Case 3	Johansen-Fisher Panel Cointegration Test: Case 4	Conclusion
rer_cpi	No	No	No	No
rer_def	No	No	No	No
rer_def_nt	No	No	No	No

APPENDIX-B

EViews Programming Code for Indonesia

```
wfopen "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-6\Indonesia.wf1"

'*****************************************************************************
'Group Plot for RER_CPI, RER_DEF, RER_DEF_NT and A_TILDE
'*****************************************************************************
group gA rer_cpi rer_def rer_def_nt A_tilde
freeze(group_plot) gA.line(x)
group_plot.setelem(1) lcolor(black) symbol(7) lpat(1)
group_plot.setelem(2) lcolor(black) symbol(4) lpat(1)
group_plot.setelem(3) lcolor(black) symbol(1) lpat(1)
group_plot.setelem(3) lcolor(black)
group_plot.options linepat
group_plot.addtext(t) Real ERs and Productivity Gap (Indonesia & U.S): 1976-2013
group_plot.addtext(b) Year
group_plot.addtext(l) RER_CPI
group_plot.addtext(l) RER_DEF
group_plot.addtext(l) RER_CPI, RER_DEF & RER_DEF_NT
group_plot.addtext(r) A_TILDE

'************************************************************
'************************************************************
create y 1976 2013
'importing data from Excel for Indonesia
import "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-6\Chapter 6.xlsx" range="Indonesia"
'****************************************************************************************************
'CASE-1: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_CPI & A_TILDE
'****************************************************************************************************
'*************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'*************************************************************
'*********************************
'Graph for Indonesia's RER_CPI
'*********************************

genr rer_cpi = rer_cpi
freeze(figure_rer_cpi) rer_cpi.line
figure_rer_cpi.addtext(t) rer_cpi (Indonesia):  1976-2013
figure_rer_cpi.addtext(b) Year
figure_rer_cpi.addtext(l) rer_cpi
figure_rer_cpi.legend(off)

'We see from the FIGURE that rer_cpi has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'****************************************************
'ADF Unit Root Test for Indonesia's RER_CPI
'****************************************************

freeze(table_6_1_1_rer_cpi_adf) rer_cpi.uroot(adf,trend,info=sic)
'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.66 which is greater than our 5% criterion -3.54.  Thus, at this point, we cannot reject the null of a unit root.
```

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_cpi_adf) rer_cpi.uroot(adf,const,trend,info=sic)
freeze(rer_cpi_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_cpi series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr rer_cpidiff = d(rer_cpi)
freeze(figure_rer_cpidiff) rer_cpidiff.line
figure_rer_cpidiff.addtext(t) drer_cpi (Indonesia):  1976-2013
figure_rer_cpidiff.addtext(b) Year
figure_rer_cpidiff.addtext(l) Drer_cpi
figure_rer_cpidiff.legend(off)
```

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

```
genr rer_cpidiff = d(rer_cpi)
freeze(table_6_1_2_rer_cpidiff1_adf) rer_cpidiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -6.78 which is now smaller than our 5% criterion -2.94. Thus, we may now reject the null of non-stationarity in first differenced series of rer_cpi. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_cpidiff1_adf) rer_cpidiff.uroot(adf,const,info=sic)
freeze(rer_cpidiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_cpi series is I(1).

'***
'DF-GLS Unit Root Test for Indonesia's RER_CPI
'***

```
freeze(table_6_1_3_rer_cpi_dfgls) rer_cpi.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.67 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of unit root.

'Now let's see if the series is difference stationary or not

```
genr rer_cpidiff = d(rer_cpi)
freeze(table_6_1_4_rer_cpidiff1_dfgls_d) rer_cpidiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -6.73 which is now smaller than our 5% criterion -1.94. Thus, we may reject the null of non-stationarity in first differenced series of rer_cpi.

''Putting it all together, I conclude that the rer_cpi series is I(1), a finding compatible with my ADF test results.

```
'**************************************************
'Graph for Indonesia's Productivity (a_tilde)
'**************************************************

genr a_tilde = a_tilde
freeze(figurea_tilde) a_tilde.line
figurea_tilde.addtext(t) a_tilde (Indonesia):  1976-2013
figurea_tilde.addtext(b) Year
figurea_tilde.addtext(l) a_tilde
figurea_tilde.legend(off)

'We see from the FIGURE that a_tilde has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'********************************************************
'ADF Unit Root Test for Indonesia's Productivity
'********************************************************

freeze(table_6_1_1_a_tilde_adf) a_tilde.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -4.46 which is smaller than our 5% criterion -3.54.  Thus, at this point, we can reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,a_tilde_adf) a_tilde.uroot(adf,const,trend,info=sic)
freeze(a_tilde_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the a_tilde series is level stationary.

'************************************************************
'DF-GLS Unit Root Test for Indonesia's Productivity
'************************************************************

freeze(table_6_1_3_a_tilde_dfgls) a_tilde.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -2.83 which is greater than our 5% criterion -3.19.  Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr a_tildediff = d(a_tilde)
freeze(table_6_1_4_a_tilde1diff1_dfgls) a_tildediff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 1.  The unit root test produces a t-value of -0.82 which is still greater than our 5% criterion -1.95. Thus, we may not reject the null of non-stationarity in first differenced series of a_tilde.

''Putting it all together, I conclude that the a_tilde series is Greater than I(1), a finding incompatible with my ADF test results.

'**********************************************
'Single Equation Cointegration Methods
'**********************************************

'**********************************************************
''Graph the suspected cointegrated series together
'**********************************************************

'The first step is to plot a graph of the suspected series.  This is very important!

group g1 rer_cpi a_tilde
freeze(figure6_1a) g1.line(x)
figure6_1a.setelem(1) lcolor(black)
```

```
figure6_1a.setelem(2) lcolor(black) lpat(8)
figure6_1a.options linepat
figure6_1a.addtext(t) rer_cpi and a_tilde (Indonesia & U.S): 1976-2013
figure6_1a.addtext(b) Year
figure6_1a.addtext(l) rer_cpi
figure6_1a.addtext(r) a_tilde

''*********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'*********************************************************

freeze(table_6_1_egc_rer_cpi) g1.coint(method=eg)

'The null hypothesis will not be rejected as suggested by sample statistics.

''********************************************
''S1.B.Error Correction Model (ECM)
'********************************************
''************************************************
'Selecting the number of lags in the VAR
'************************************************

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var table_6_1_var1.ls 1 4   g1
freeze(table_6_1_var1_lagtest1) table_6_1_var1.laglen(4)
freeze(table_6_1_var1_lagtest2) table_6_1_var1.testlags

'The lag length test above indicates that the VAR has 3 lags.

var table_6_1_var2.ls 1 3  g1
freeze(table_6_1_var2_artest1) table_6_1_var2.correl
freeze(table_6_1_var2_artest2) table_6_1_var2.qstats(12)
freeze(table_6_1_var2_artest3) table_6_1_var2.arlm(12)

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_cpi c a_tilde
genr ec1 = resid


var table_6_1_eg2a_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3))

var table_6_1_eg2b_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3)) d(a_tilde(-1))

var table_6_1_eg2c_cpi.ls 0 0 d(rer_cpi)   @  c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.

var table_6_1_eg2a_cpi.ls 0 0 d(rer_cpi)   @   c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3))
freeze(table_6_1_eg2a1_cpi_artest1) table_6_1_eg2a_cpi.correl
freeze(table_6_1_eg2a2_cpi_artest2) table_6_1_eg2a_cpi.qstats(12)
freeze(table_6_1_eg2a3_cpi_artest3) table_6_1_eg2a_cpi.arlm(12)

'The residuals are absolutely white noise.

''*************************
''Estimating EC Model
'*************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:
```

```
equation table_6_1_ecm_rer_cpi.ls(n) d(rer_cpi) c ec1(-1) d(rer_cpi(-1)) d(rer_cpi(-2)) d(rer_cpi(-3))

'Note that the SR effect is significant as the error correction coefficient -0.14 is statistically significant at better than 1% significance level.

''************************************************
''S2.A & S2.B: Obtaining LR Coefficients
'*************************************************
'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Indonesia against U.S.

equation table_6_1_LReqn1a_fmols.cointreg(method=fmols) rer_cpi a_tilde

equation table_6_1_LReqn1b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2 ) rer_cpi a_tilde

'The BS coefficient obtained through FMOLS and DOLS estimator are 0.41 and 0.37 but statistically insignificant. Thus, there is 'NO' evidence in support of BS effect existing for Indonesia.

''********************************************
'Multivariate Cointegration Approach
'********************************************

''**********************************************************
''Check if the VAR (2) model is dynamically stable
'***********************************************************
freeze (table_6_1_var2_varstable) table_6_1_var2.arroots(graph)

'The model is dynamically stable.

''*************************************************************************
''M1.A & M1.B: Identifying the number of cointegrating vectors
'**************************************************************************

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_6_1_var2_coint1) table_6_1_var2.coint(s,3)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. All the results indicate 0 cointegrating vectors.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).
```

```
''*****************************************************************
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'*****************************************************************

'For estimating the LR relationship, corresponding VEC command is:

var table_6_1_vec1d.ec(d,1) 1 2 rer_cpi a_tilde

'CONCLUSION:  I conclude that rer_cpi and a_tilde are not cointegrated in the Indonesia's data.

'************************************************************************************************
'CASE-2: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_DEF & A_TILDE
'************************************************************************************************
'***********************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'***********************************************************
'*************************************
'Graph for Indonesia's RER_DEF
'*************************************

genr rer_def = rer_def
freeze(figure_rer_def) rer_def.line
figure_rer_def.addtext(t) rer_def (Indonesia):  1976-2013
figure_rer_def.addtext(b) Year
figure_rer_def.addtext(l) rer_def
figure_rer_def.legend(off)

'We see from the FIGURE that rer_def has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'***************************************************
'ADF Unit Root Test for Indonesia's RER_DEF
'***************************************************

freeze(table_6_1_1_rer_def_adf) rer_def.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.54 which is greater than our 5% criterion -3.54.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_adf) rer_def.uroot(adf,const,trend,info=sic)
freeze(rer_def_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

genr rer_defdiff = d(rer_def)
freeze(figure_rer_defdiff) rer_defdiff.line
figure_rer_defdiff.addtext(t) drer_def (Indonesia):  1976-2013
figure_rer_defdiff.addtext(b) Year
figure_rer_defdiff.addtext(l) drer_def
figure_rer_defdiff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr rer_defdiff = d(rer_def)
freeze(table_6_1_2_rer_defdiff1_adf) rer_defdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -6.41 which is now smaller than our 5% criterion -2.94. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_defdiff1_adf) rer_defdiff.uroot(adf,const,info=sic)
freeze(rer_defdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def series is I(1).

'**
'DF-GLS Unit Root Test for Indonesia's RER_DEF
'**

```
freeze(table_6_1_3_rer_def_dfgls) rer_def.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.69 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of unit root.

'Now let's see if the series is difference stationary or not

```
genr rer_cpidiff = d(rer_def)
freeze(table_6_1_4_rer_defdiff1_dfgls_d) rer_defdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -6.37 which is now smaller than our 5% criterion -1.95. Thus, we may reject the null of non-stationarity in first differenced series of rer_def.

''Putting it all together, I conclude that the rer_def series is I(1), a finding compatible with my ADF test results.

'**
'Single Equation Cointegration Methods
'**
'***
''Graph the suspected cointegrated series together
'***

'The first step is to plot a graph of the suspected series. This is very important!

```
group g2 rer_def a_tilde
freeze(figure6_1b) g2.line(x)
figure6_1b.setelem(1) lcolor(black)
figure6_1b.setelem(2) lcolor(black) lpat(8)
figure6_1b.options linepat
figure6_1b.addtext(t) rer_def and a_tilde (Indonesia & U.S): 1976-2013
figure6_1b.addtext(b) Year
figure6_1b.addtext(l) rer_def
figure6_1b.addtext(r) a_tilde
```

''***
''S1.A.Engle-Granger Approach to Cointegration
'***

```
freeze(table_6_1_egc_rer_def) g2.coint(method=eg)
```

'The null hypothesis will not be rejected as suggested by sample statistics.

"***
"S1.B.Error Correction Model (ECM)
'***

"***
'Selecting the number of lags in the VAR
'***

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

```
var table_6_1_var3.ls 1 4   g2
freeze(table_6_1_var3_lagtest1) table_6_1_var3.laglen(4)
freeze(tale_6_1_var3_lagtest2) table_6_1_var3.testlags
```

'The lag length test above indicates that the VAR has 3 lags.

```
var table_6_1_var4.ls 1 3  g2
freeze(table_6_1_var4_artest1) table_6_1_var4.correl
freeze(table_6_1_var4_artest2) table_6_1_var4.qstats(12)
freeze(table_6_1_var4_artest3) table_6_1_var4.arlm(12)
```

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

```
genr resid = 0
equation eg.ls rer_def c a_tilde
genr ec2 = resid

var table_6_1_eg2a_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3))

var table_6_1_eg2b_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(a_tilde(-1))

var table_6_1_eg2c_def.ls 0 0 d(rer_def)   @  c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(a_tilde(-1)) d(a_tilde(-2))
```

'The evidence suggests that Model C is best. Now we test that model for serial correlation.

```
var table_6_1_eg2a_def.ls 0 0 d(rer_def)   @   c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(a_tilde(-1)) d(a_tilde(-2))
freeze(table_6_1_eg2a1_def_artest1) table_6_1_eg2a_def.correl
freeze(table_6_1_eg2a2_def_artest2) table_6_1_eg2a_def.qstats(12)
freeze(table_6_1_eg2a3_def_artest3) table_6_1_eg2a_def.arlm(12)
```

'The residuals are absolutely white noise.

"*************************
'Estimating EC Model
'*************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

```
equation table_6_1_ecm_rer_def.ls(n) d(rer_def) c ec2(-1) d(rer_def(-1)) d(rer_def(-2)) d(rer_def(-3)) d(a_tilde(-1)) d(a_tilde(-2))
```

'Note that the SR effect is significant as the error correction coefficient -0.11 is statistically significant at better than 1% significance level.
"***
"S2.A & S2.B: Obtaining LR Coefficients
'***

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient, i.e., BS coefficient for Indonesia against U.S.

```
equation table_6_1_LReqn2a_fmols.cointreg(method=fmols) rer_def a_tilde

equation table_6_1_LReqn2b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2 ) rer_def a_tilde

'The BS coefficients obtained through FMOLS and DOLS estimators are 1.62 and 1.84 respectively. The two coefficient are positive and are statistically significant. Thus, there is sufficient evidence in support of valid BS effect existing for Indonesia.

''******************************************
'Multivariate Cointegration Approach
'******************************************

''*********************************************************
''Check if the VAR (4) model is dynamically stable
'*********************************************************
freeze(var4_varstable) table_6_1_var4.arroots(graph)

'The model is dynamically stable.

''**********************************************************************
''M1.A & M1.B: Identifying the number of cointegrating vectors
'**********************************************************************

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_6_1_var4_coint2) table_6_1_var4.coint(s,3)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. Trace statistic of Case 3 indicate 1 cointegrating vector.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

''******************************************************************
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'******************************************************************

' For estimating the LR relationship, corresponding VEC command is:

var table_6_1_vec2d.ec(d,1) 1 2 rer_def a_tilde

'CONCLUSION:  I conclude that rer_def and a_tilde are not cointegrated in the Indonesia's data.

'**********************************************************************************************************
'CASE-3: ESTIMATING BALASSA-SAMUELSON EFFECT FOR RER_DEF_NT & A_TILDE
'**********************************************************************************************************
'************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'************************************************************
'******************************************
'Graph for Indonesia's RER_DEF_NT
'******************************************

genr rer_def_nt = rer_def_nt
freeze(figure_rer_def_nt) rer_def_nt.line
figure_rer_def_nt.addtext(t) rer_def_nt (Indonesia):  1976-2013
figure_rer_def_nt.addtext(b) Year
figure_rer_def_nt.addtext(l) rer_def_nt
figure_rer_def_nt.legend(off)
```

'We see from the FIGURE that rer_def_nt has time trend to it. So we would include both an intercept and a time trend in our unit root regression equations.

'**
'ADF Unit Root Test for Indonesia's RER_DEF_NT
'**

freeze(table_6_1_1_rer_def_nt_adf) rer_def_nt.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -0.99 which is greater than our 5% criterion -3.54. Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_nt_adf) rer_def_nt.uroot(adf,const,trend,info=sic)
freeze(rer_def_nt_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_nt series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

genr rer_def_ntdiff = d(rer_def_nt)
freeze(figure_rer_def_ntdiff) rer_def_ntdiff.line
figure_rer_def_ntdiff.addtext(t) drer_def_nt (Indonesia): 1976-2013
figure_rer_def_ntdiff.addtext(b) Year
figure_rer_def_ntdiff.addtext(l) drer_def_nt
figure_rer_def_ntdiff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_6_1_2_rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -6.19 which is now smaller than our 5% criterion -2.94. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def_nt. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)
freeze(rer_def_ntdiff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_nt series is I(1).

'***
'DF-GLS Unit Root Test for Indonesia's RER_DEF_NT
'***

freeze(table_6_1_3_rer_def_nt_dfgls) rer_def_nt.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.14 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not
Genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_6_1_4_rer_def_nt1diff1_dfgls) rer_def_ntdiff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -6.15 which is now smaller than our 5% criterion -1.94. Thus, we may reject the null of non-stationarity in first differenced series of rer_def_nt.

```
''Putting it all together, I conclude that the rer_def_nt series is I(1), a finding compatible with my ADF test results.

'************************************************
'Single Equation Cointegration Methods
'************************************************

'*************************************************************
''Graph the suspected cointegrated series together
'*************************************************************

'The first step is to plot a graph of the suspected series.  This is very important!

group g3 rer_def_nt a_tilde
freeze(figure6_1) g3.line(x)
figure6_1.setelem(1) lcolor(black)
figure6_1.setelem(2) lcolor(black) lpat(8)
figure6_1.options linepat
figure6_1.addtext(t) rer_def_nt and a_tilde (Indonesia & U.S): 1976-2013
figure6_1.addtext(b) Year
figure6_1.addtext(l) rer_def_nt
figure6_1.addtext(r) a_tilde

''**********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'**********************************************************

freeze(table_6_1_egc_rer_def_nt) g3.coint(method=eg)

'The null hypothesis will not be rejected as suggested by sample statistics.
''*********************************************
''S1.B.Error Correction Model (ECM)
'*********************************************
''************************************************
'Selecting the number of lags in the VAR
'*************************************************

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var table_6_1_var5.ls 1 6   g3
freeze(table_6_1_var5_lagtest1) table_6_1_var5.laglen(4)
freeze(tale_6_1_var5_lagtest2) table_6_1_var5.testlags

'The lag length test above indicates that the VAR has 3 lags.

var table_6_1_var6.ls 1 3  g3
freeze(table_6_1_var6_artest1) table_6_1_var6.correl
freeze(table_6_1_var6_artest2) table_6_1_var6.qstats(12)
freeze(table_6_1_var6_artest3) table_6_1_var6.arlm(12)

'We now try different lags of d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_def_nt c a_tilde
genr ec3 = resid

var table_6_1_eg2a_def_nt.ls 0 0 d(rer_def_nt)   @  c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3))

var table_6_1_eg2b_def_nt.ls 0 0 d(rer_def_nt)   @  c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(a_tilde(-1))

var table_6_1_eg2c_def_nt.ls 0 0 d(rer_def_nt)   @  c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.
```

var table_6_1_eg2a_def_nt.ls 0 0 d(rer_def_nt) @ c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3))
freeze(table_6_1_eg2a1_def_nt_artest1) table_6_1_eg2a_def_nt.correl
freeze(table_6_1_eg2a2_def_nt_artest2) table_6_1_eg2a_def_nt.qstats(12)
freeze(table_6_1_eg2a3_def_nt_artest3) table_6_1_eg2a_def_nt.arlm(12)

'The residuals are absolutely white noise.

''*************************
'Estimating EC Model
'**************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_6_1_ecm_rer_def_nt.ls(n) d(rer_def_nt) c ec3(-1) d(rer_def_nt(-1)) d(rer_def_nt(-2)) d(rer_def_nt(-3))

'Note that the SR effect is significant as the error correction coefficient -0.13 is statistically significant at better than 1% significance level.

''***
''S2.A & S2.B: Obtaining LR Coefficients
'**

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Indonesia against U.S.

equation table_6_1_LReqn3a_fmols.cointreg(method=fmols) rer_def_nt a_tilde

equation table_6_1_LReqn3b_dols.cointreg(method=dols, trend=constant, lag=2,lead=2) rer_def_nt a_tilde

'The BS coefficients obtained through FMOLS and DOLS estimators are -0.23 and 0.28. The two coefficients are negative and/or statistically insignificant. Thus, there is insufficient evidence in support of BS effect existing for Indonesia.

''**
'Multivariate Cointegration Approach
'**

''**
''Check if the VAR (6) model is dynamically stable
'***
freeze(var6_varstable) table_6_1_var6.arroots(graph)

'The model is dynamically stable.

''**
''M1.A & M1.B: Identifying the number of cointegrating vectors
'***

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_6_1_var6_coint3) table_6_1_var6.coint(s,3)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. Trace statistic of Case 3 indicate 1 cointegrating vector.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

```
''*******************************************************************
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'********************************************************************

' For estimating the LR relationship, corresponding VEC command is:

var table_6_1_vec3d.ec(d,1) 1 2 rer_def_nt a_tilde

'CONCLUSION:  I conclude that rer_def_nt and a_tilde are not cointegrated in the Indonesia's data.
```

CHAPTER 7: AN INVESTIGATION OF DOMESTIC VERSION OF THE BALASSA-SAMUELSON HYPOTHESIS

In Chapters Five and Six, I did not find any sizeable evidence in support of the Balassa-Samuelson (BS) effect for developing Asian countries. I empirically tested the model using two alternative data sets, offering narrow as well as broad sectoral divisions, through both time series and panel data estimators. As a robustness check, I employed three alternative price deflators to construct different versions of real exchange rate series. Still, I obtained insufficient support for the hypothesis, suggesting the notable absence of the BS effect for the subject economies.

According to the BS hypothesis, productivity improvements in the tradable sector of a country should cause prices in the nontradable sector of that country to rise relative to tradable prices. This is referred to as the "domestic" version of the BS effect. My failure to find evidence of the BS hypothesis in international prices has turned me to examine intra-country prices to see whether any evidence of a BS effect can be found at that level.

A sizeable number of studies have tested this domestic version of the BS hypothesis (see Canzoneri et al., 1999; Egert, 2002, 2005; Lojschova, 2003; Mihaljek and Klau, 2004; Lee and Tang, 2007; Funda et al., 2008; Thomas and King, 2008). These studies investigate the domestic version of the model as key a driver of the standard BS mechanism. In the literature, the domestic version of the BS hypothesis is known as the Baumol-Bowen effect. Baumol and Bowen (1966) argue that within a country, there is a broad tendency of the prices of service-intensive industries (education, health care, banking, etc.) to rise over time as, historically, productivity growth in these sectors has tended to be slower than in capital-intensive, manufacturing industries.

To investigate whether the “domestic” price mechanism of the BS model holds, I will conduct time-series and panel data empirical tests. However, to set up those tests, I first lay out the associated theoretical foundation.

7.1 Long-run Association between Domestic Relative Prices and Productivities

From Chapter Two, the long-run association between relative sectoral prices and productivities is recalled to be:

(7.1) $p_t^{NT} - p_t^{T} = \frac{\gamma}{\delta}(a_t^{T} - a_t^{NT}) + c$,

where $(p_t^{NT} - p_t^{T})$ is the relative sectoral price of nontradables, $(a_t^{T} - a_t^{NT})$ is the relative sectoral productivity of tradables, and $\left(\frac{\gamma}{\delta}\right)$ is the sectoral labour intensity ratio. Note that $\left(\frac{\gamma}{\delta}\right) > 0$. This implies that relative productivity gains in the traded sector will cause the relative sectoral price of nontradables to rise. This should induce a positive long-run correlation between the respective relative productivity and relative price variables. From equation (7.1), I obtain the following estimable form of the domestic version of the BS effect:

(7.2) $p_t^{NT} - p_t^{T} = \theta_0 + \theta_1 a_t + \varepsilon_{1t}$

where $a_t = (a_t^{T} - a_t^{NT})$. If the *domestic* BS effect holds true then θ_1, the slope coefficient on the relative sectoral productivity of tradables, should be positive and statistically significant. Unlike preceding chapters, where the real exchange rate was constructed by using three alternative measures of prices, this chapter involves only one measure of the internal real exchange rate, i.e., the GDP deflator based sectoral price ratio of nontradables to tradables. Detailed notes on sectoral division for each country, construction of country-specific relative sectoral prices and productivity and their data sources can be found in Chapter Three.

The assumed value of ${}^{\gamma}/_{\delta}$ plays an important role in testing the domestic version of the BS model. Thomas and King (2008) establish an equi-proportionate relationship between biased sectoral productivity and relative sectoral prices. Others, such as Mihaljik and Klau (2004), allow for equi-proportionate changes only if both sectors have the same degree of labour intensity, i.e., $\gamma = \delta$. In my analysis, the valid existence of a domestic BS effect is not conditional upon the equivalence of ${}^{\gamma}/_{\delta} = 1$. In general, it seems plausible that the degree of labour intensity may vary from industry to industry. Thus, in my tests, I allow for disproportionate labour shares across sectors. This approach is consistent with the empirical findings of some earlier studies, confirming the validity of a domestic BS effect, but with disproportionate effects of productivity differentials on the internal price ratio (Mihaljik and Klau (2004); Egert, 2005; Lee and Tang, 2007; Funda et al., 2008; Thomas and King, 2008).

7.2 Country Study

In this section, I estimate the domestic version of the BS effect country by country, using the same time-series estimation techniques I have used in earlier chapters. The country results will be concluded to support the existence of a valid domestic BS effect whenever the imbalanced productivity patterns of traded and nontraded sectors are positively and significantly associated with appreciation in the relative price of nontradables in that country. The domestic version of the model is tested using equation (7.2) above. Readers may refer to the EViews program code for Indonesia, provided in the Appendix at the end of this chapter, to confirm the appropriateness of the econometric procedures I use.

7.2.1 Indonesia

As always, I start with Indonesia. The time plots of internal relative prices of nontradables and relative sectoral productivity can be seen in FIGURE 7.1. Over the time, the two series are generally co-moving downwards. From the visual inspection, I see evidence for a causal effect between country sectoral productivity bias and trend appreciation in relative prices of nontradables. The price series appear to be adjusting with modest speed to induce self-corrections, causing the internal real exchange rate of the country to return to long-run equilibrium.

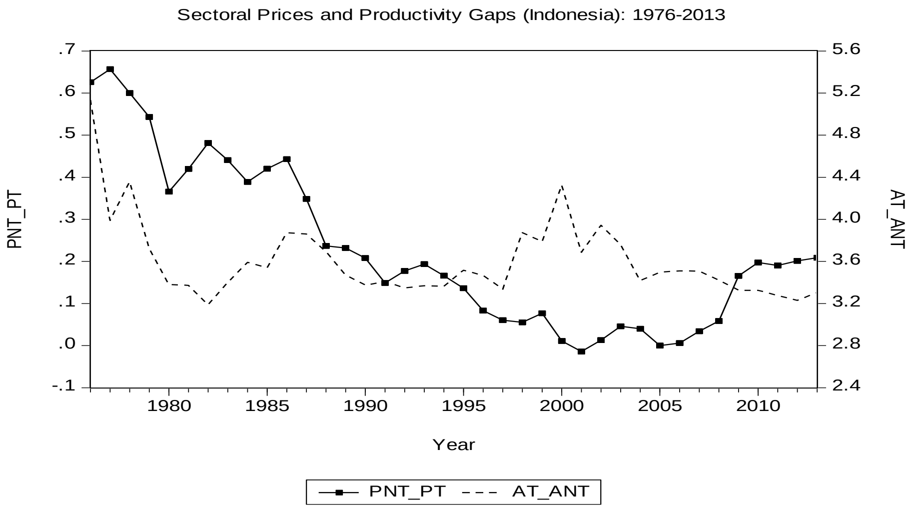
In TABLE 7.1, the formal econometric tests for deciding the order of integration of model variables unanimously state that the price series is integrated of order one. However, for the productivity ratio, the evidence is mixed. According to the ADF unit root test, the series is unit root in levels; whereas the DF-GLS unit root test suggests that the order of integration is greater than one. These results leave me indecisive on the actual order of integration of the productivity series.

Turning now to the single equation cointegration test based on the regression residuals, the EG test is unable to reject the null hypothesis of no cointegration. The tau-values are tested against MacKinnon (1996)'s critical values, indicating no long-run causality between the productivity and price differential series for Indonesia. However, the Error Correction Model (ECM) results challenge the EG test findings. The ECM yields a negative and statistically significant Error Correction (EC) coefficient value of -0.09. This implies that in each year, 9 percent of the total deviation in the price series from long-run equilibrium is corrected/adjusted by the series itself. I next estimate the FMOLS and DOLS cointegration regressions to obtain the long-run BS estimates for the country. Both tests do not support the BS hypothesis, as the two long-run BS coefficients are statistically insignificant. Thus, the single equation cointegration approach does not find evidence of a long-run association between Indonesia's internal relative prices and relative productivity of the traded and nontraded sectors.

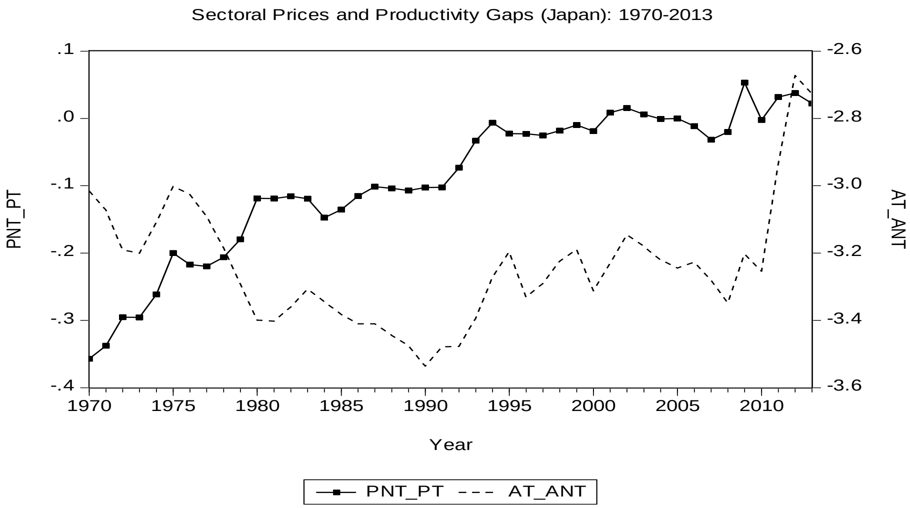
Similar to the single equation cointegration tests results, there is some support in favour of long-run co-movement between model variables using the multivariate cointegration method. The Trace and Maximum Eigenvalue statistics of Case 4 of the Johansen ML test both indicate the presence of one valid cointegrating vector. This allows me to estimate short-run to long-run dynamics of the model using the Vector Error Correction Model (VECM) specification.

The VECM estimates do not support the valid existence of a causal effect between the sectoral productivity and sectoral price ratios. Although, the short-run to long-run adjustment coefficient (EC coefficient) is negative and statistically significant, the long-run BS coefficient is wrong-signed, and statistically significant. In addition, the model does not meet the condition of weak exogeneity, as the EC_a coefficient is statistically significant.

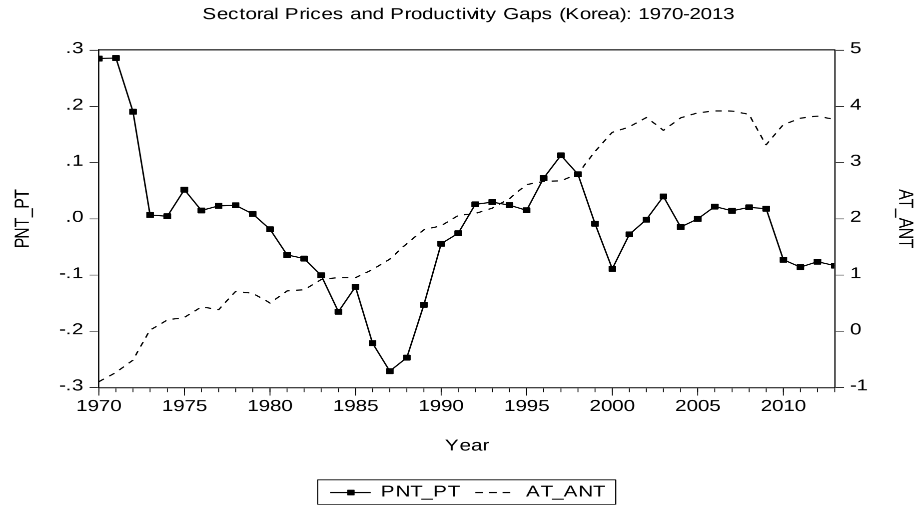
FIGURE 7.1: Plots for Domestic Sectoral Prices and Productivity Differentials



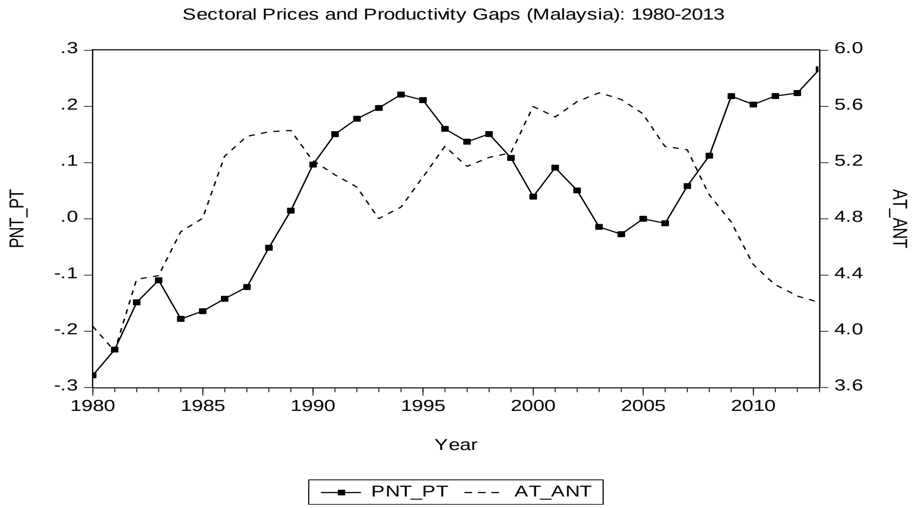
Indonesia



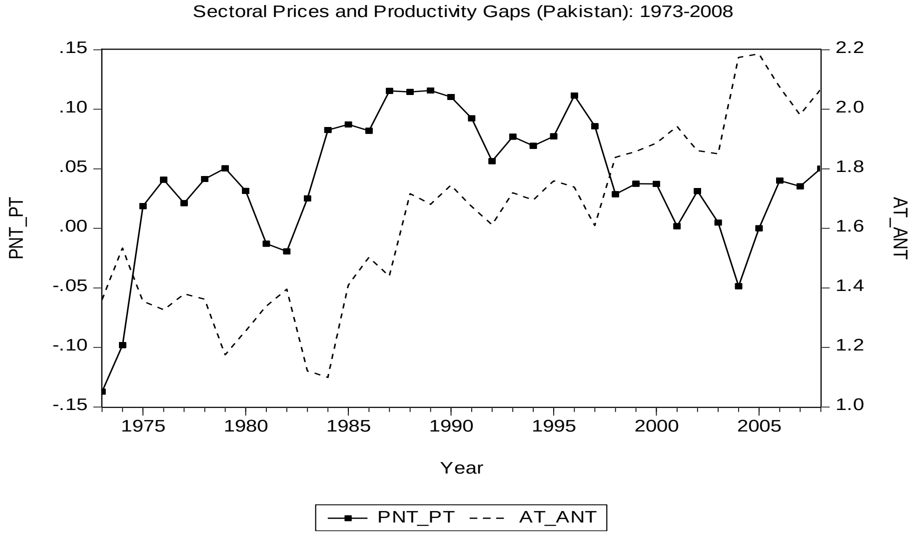
Japan



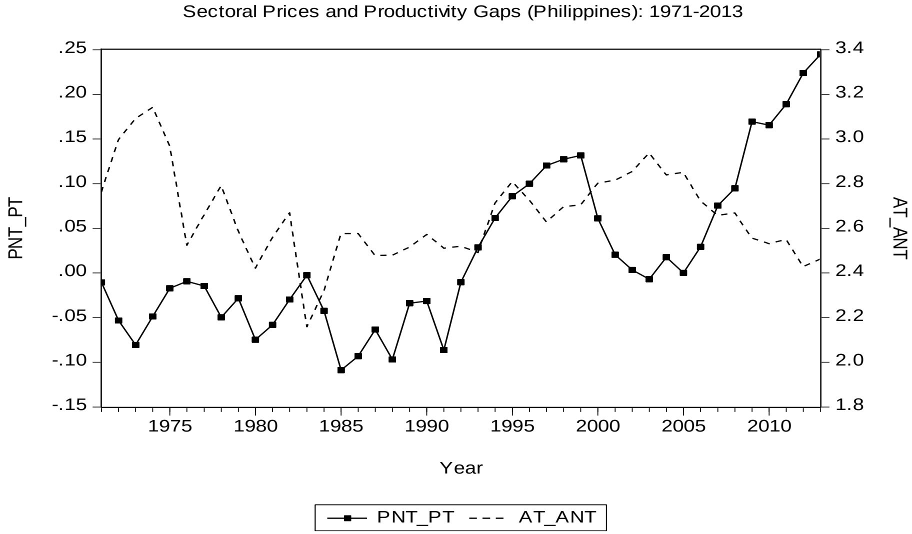
Korea



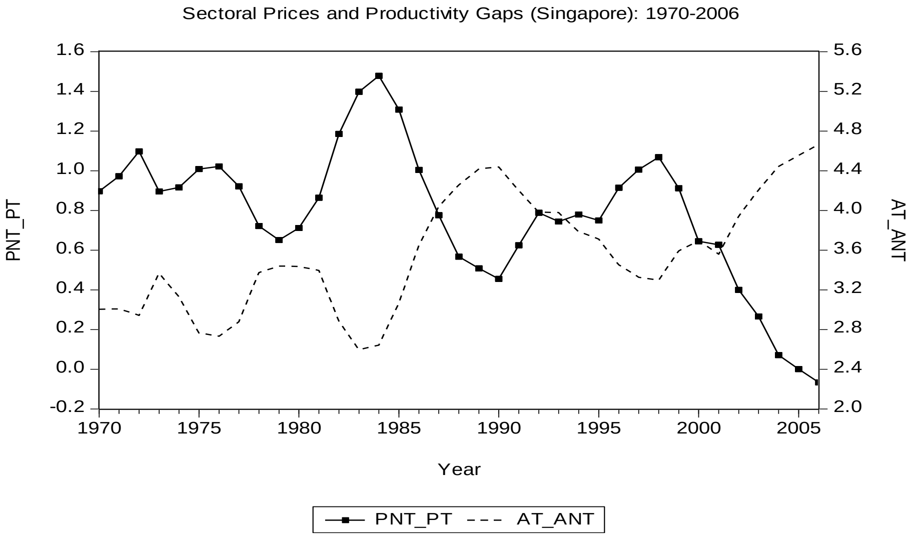
Malaysia



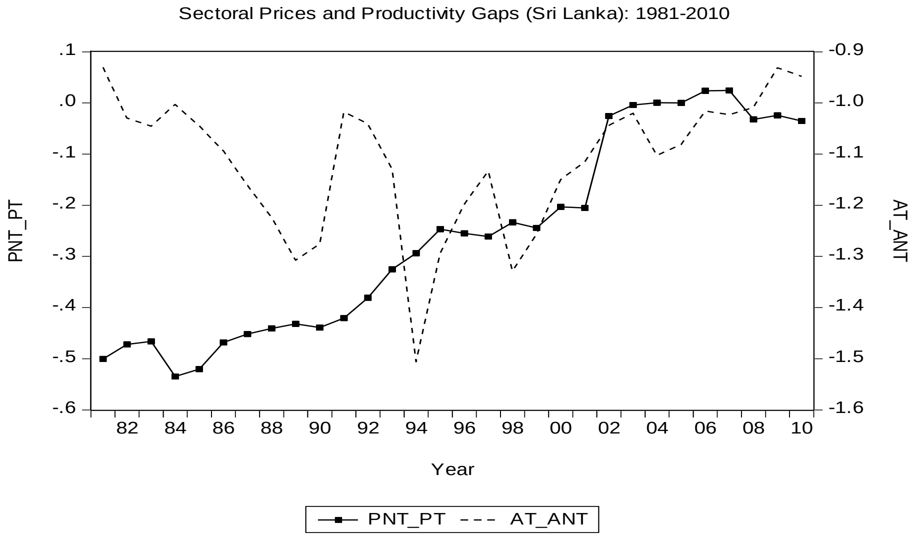
Pakistan



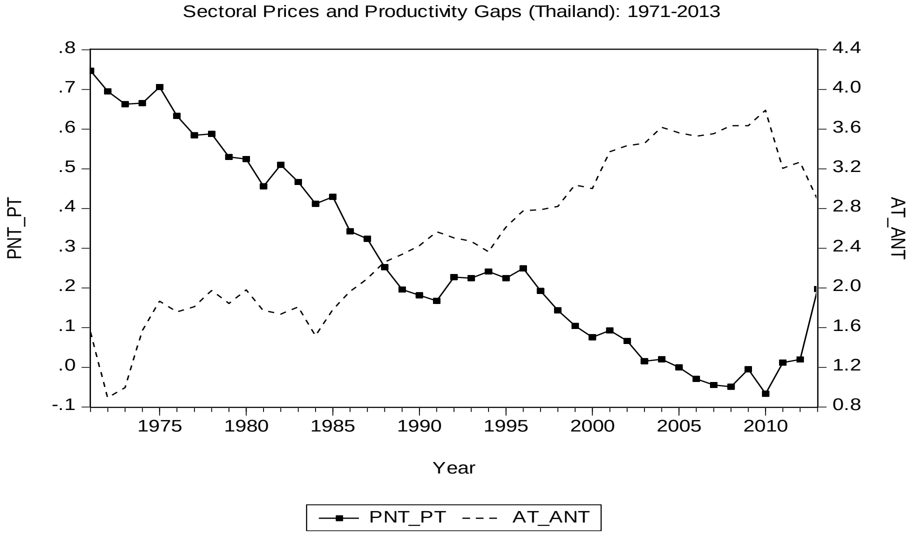
Philippines



Singapore



Sri Lanka


Sectoral Prices and Productivity Gaps (Thailand): 1971-2013
PNT_PT
AT_ANT
.8
.7
.6
.5
.4
.3
.2
.1
.0
-.1
4.4
4.0
3.6
3.2
2.8
2.4
2.0
1.6
1.2
0.8
1975
1980
1985
1990
1995
2000
2005
2010
Year
PNT_PT
AT_ANT

Thailand

TABLE 7.1: Cointegration Tests Results for Indonesia (1976-2013)[90]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	I(0)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[91]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient[92]	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
$p^{NT} - p^{T}$	No	3	-0.09 [-1.97]	0.15 [0.80]	-0.04 [-0.11]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	2	2	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	1	1	See below

Vector Error Correction Model

	Lags	White Noise Residuals	$EC_{p^{NT}-p^{T}}$	EC_{a}	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$p^{NT} - p^{T}$	2	Yes	-0.10 [-3.71]	-0.34 [-2.65]	-0.95 [-3.91]	No

[90] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[91] t-values are given in squared-brackets.
[92] The test regression also contain first and second lagged-differences of a.

Taken together, my empirical analysis does not find evidence to support the BS hypothesis for Indonesia.

7.2.2 Japan

A visual inspection of the time-series plots of the relative prices of nontradables and the sectoral productivity differential for Japan does not produce much support for a domestic BS effect. The two series trend in dissimilar directions, accompanied by infrequent intersections. These behaviours make a long-run association between the variables doubtful.

The empirical estimates for Japan are reported in TABLE 7.2 below. Overall, there are no traces of a BS effect being operative with respect to sectoral prices and productivities. Though the model variables are unit root in levels, there is no evidence from the EG residuals test that the series cointegrate. In contrast, the cointegration test based on the error correction coefficient in the ECM produces a different conclusion. It finds that the series are cointegrated, with a speed of adjustment to long-run equilibrium of 10 percent every year. However, when I estimate the long-run relationship between these variables using FMOLS and DOLS, I find that the respective BS coefficient is statistically insignificant. I conclude that the single equation cointegration approach does not produce any evidence of a BS effect for Japan.

I get even less far when I use the multivariate cointegration procedure. The Trace and Maximum Eigenvalues statistics for the Johansen ML cointegration test (under both Case 3 and Case 4) find that the series do not co-move, as no evidence for a long-run relationship between the variables can be found. Both model specifications produce a rank of zero. Thus, through both single equation and multivariate cointegration methods, I conclude that the domestic BS effect does not hold valid for Japan.

TABLE 7.2: Cointegration Tests Results for Japan (1970-2013)[93]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[94]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
$p^{NT} - p^{T}$	No	1	-0.10 [-4.90]	-0.09 [-0.66]	-0.16 [-0.78]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

93 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
94 t-values are given in squared-brackets.

7.2.3 Korea

The two time-series (relative prices of nontradables and sectoral productivity gap for Korea) are plotted in FIGURE 7.1. For most of the sample period, the relative price of nontradables shows no relationship with the steadily increasing productivity gap between the tradable and nontradable sectors. Their occasional intersection does not provide visual support for a cointegrating relationship.

The empirical estimates for Korea are reported in TABLE 7.3. The EG single equation residuals test for cointegration detects a valid long-run relationship between the relative prices of nontradables and the sectoral productivity ratio. The test results reveal a statistically significant long-run co-movement on the part of prices and its regressor, as evident from the tau-statistic rejecting the null hypothesis of no cointegration at better than a five percent significance level. This stands in contrast to the ECM test results, which indicate a lack of long-run co-movement between the variables. The statistically insignificant EC coefficient suggests that short-termed fluctuations in the price series are not corrected by movements in the series towards a long-run relationship.

Because of the positive test results from the EG residuals test, I proceed with estimating long-run relationship between the model variables using the FMOLS and DOLS estimators. The estimated BS coefficients are statistically insignificant. These results do not support the existence of a long-run BS effect for the country.

.

TABLE 7.3: Cointegration Tests Results for Korea (1970-2013)[95]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(0)**	I(1)***	Inconclusive
a	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[96]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient[97]	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	Yes	3	-0.10 [-1.48]	-0.01 [-0.42]	0.01 [0.75]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

95 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
96 t-values are given in squared-brackets.
97 The test regression also contain first and second lagged-differences of a.

Multivariate cointegration tests also do not detect evidence of a long-run association between model variables. I obtain ranks of zero for both Trace and Max Eigenvalue statistics under both specifications of the Johansen ML cointegration test. This indicates that the two series do not have a tendency to return to each other over time. As a result, I do not proceed further to estimate a cointegrating equation. The single equation and multivariate cointegration approaches are consistent in their lack of evidence of a BS effect. I thus conclude that the BS hypothesis does not hold for Korea.

7.2.4 Malaysia

In FIGURE 7.1, the time plots of the model variables for Malaysia are given. From the visual assessment, it seems like there is some chance of a long-run association between the two series. For almost the entire period, the two series are co-moving in a common direction (before 2007 in specific). Furthermore, their regular intersection points to the potential of establishing long-run cointegration. It is only at the end of the sample period that there is uncertainty about the two series returning to each other.

The single equation cointegration test results generally support the visual assessment of FIGURE 7.1 when it comes to the existence of a long-run relationship. The error correction coefficient in the ECM find evidence to support the existence of cointegration between the two variables. However, this long-run association does not hold in the desired way. The two long-run cointegration regression estimators, i.e., the FMOLS and DOLS, each produce statistically insignificant long-run BS coefficient. Thus, the BS effect does not seem to hold for Malaysia when the hypothesis is tested using the single equation cointegration models.

TABLE 7.4: Cointegration Tests Results for Malaysia (1980-2013)[98]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	Greater than I(1)	I(1)**	Inconclusive

Single Equation Cointegration Approach[99]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	No	1	-0.09 [-2.13]	0.02 [0.35]	0.06 [0.68]	No

Multivariate Cointegration Approach

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
$p^{NT} - p^{T}$	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
$p^{NT} - p^{T}$	1	1	See below

Vector Error Correction Model

	Lags	White Noise Residuals	$EC_{p^{NT}-p^{T}}$	EC_{a}	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
$p^{NT} - p^{T}$	0	Yes	-0.08 [-1.82]	-0.48 [-2.28]	-0.16 [-1.78]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$p^{NT} - p^{T}$	0	Yes	-0.03 [-1.40]	-0.27 [-2.97]	-0.25 [-1.47]	No

98 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
99 t-values are given in squared-brackets.

The results generated by the multivariate cointegration test are consistent with the single equation results. Both specifications of the Johansen ML test reveal the existence of one valid cointegrating vector. Both the Trace and Maximum Eigenvalue statistics indicate a rank of one. As a result, I proceed by estimating the VECM to obtain estimates of the speed of adjustment parameters and the cointegrating equation.

The two specifications of the VEC model (Cases 3 and 4) both reject the BS hypothesis. The three pre-conditions for the BS hypothesis are not all met. For Case 4, short-run deviations in the cointegrating relationship do not result in adjustments in the relative price of nontradables. For both cases, productivity turns out to be significantly endogenous as short-run fluctuations in relative prices are being significantly corrected by the sectoral productivity gap. And, finally, the long-run BS coefficient is negative and/or statistically insignificant. I conclude that there is no evidence to support the BS hypothesis for Malaysia.

7.2.5 Pakistan

From the visual inspection of time plots of relative prices of nontradables and productivity differential of Pakistan, there is a lack of support for the domestic BS effect. Though, the two series are trending in similar directions for most of the sample period, they only infrequently intersect.

The EG and the ECM based single equation cointegration tests both suggest a long-run co-movement between the model variables. The tau-statistics, produced under the EG test, reject the null of no cointegration at the five percent statistical significance. This finding is supported by the ECM as the EC coefficient is estimated to be -0.27 and statistically significant.

TABLE 7.5: Cointegration Tests Results for Pakistan (1973-2008)[100]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(0)**	I(1)**	Inconclusive
a	`Yes	I(1)***	I(0)**	Inconclusive

Single Equation Cointegration Approach[101]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient[102]	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	Yes	1	-0.27 [-3.74]	-0.01 [-0.23]	-0.04 [-0.82]	No

`Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

[100] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[101] t-values are given in squared-brackets.
[102] The test regression also contain first and second lagged-differences of a.

However, estimation of the long-run relationship using FMOLS and DOLS does not support the BS hypothesis. Both coefficients are negative and statistically insignificant. Thus no evidence of a BS effect is found when I use the single equation cointegration model for Pakistan.

This conclusion is supported when I turn to the multivariate cointegration approach, only now I don't even find evidence that a long-run relationship exists. The Trace and Maximum Eigenvalue statistics for both Cases 3 and 4 produce the conclusion that the rank of the system is zero. This indicates that the series do not share a long-run relationship. This of course argues against the BS hypothesis, since it indicates that there is no relationship between the productivity gap and the relative price ratio for nontradables. As a result, I once again conclude that the BS effect does not hold.

7.2.6 Philippines

The visual inspection of relative prices (nontradables) and productivity ratio of Philippines from FIGURE 7.1, provides some support for a long-run association between two variables, owing to their regular intersection. The two series trend in a common direction (more visibly before the year 2000) for most of the sample period. Also, the speed of adjustment contributed by the price ratio to correct its short-run deviations seems to be relatively high.

Turning now to the single equation cointegration test results of TABLE 7.6, I find no sign of long-run cointegration between the relative price ratio of nontradables and the productivity gap. In this, both the EG residuals test and the ECM model are consistent. The two tests fail to reject the null of no cointegration with acceptable statistical significance. The results indicate the absence of a long-run association between internal relative sectoral prices and productivity, thus failing to support the BS hypothesis for the Philippines.

TABLE 7.6: Cointegration Tests Results for Philippines (1971-2013)[103]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[104]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	No	4	-0.02 [-0.17]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

[103] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[104] t-values are given in squared-brackets.

The results from the multivariate cointegration approach are in agreement. The Trace and Maximum Eigenvalue statistics under both cases of the Johansen ML cointegration test (Case 3 and Case 4) produce a rank of zero for the estimated model. This argues against the existence of a long-run association between productivity differentials and relative price movements. I thus conclude that a domestic BS effect does not hold for the Philippines.

7.2.7 Singapore

For Singapore, the model series for estimating the domestic BS hypothesis are given in FIGURE 7.1. The price and productivity series are found to be co-moving in a common direction for almost the entire data period. The fact that they only infrequently intersect argues against there being a long-run, integrating relationship. On the other hand, the aggressive movements in internal relative prices during the middle period of the sample indicate the possibility of adjustments to short-run deviations, so that a stable internal real exchange rate may be maintained in the long-run.

The empirical estimates for Singapore are reported in TABLE 7.7 below. Overall, there are no traces of BS effect existing for the country, according to the single equation cointegration tests estimates. The estimated tau-statistics of the EG model cannot reject the null hypothesis of no cointegration at the ten percent significance level. The same holds true for the ECM, where the test yields a statistically insignificant EC coefficient. Thus there is no evidence in favour of the BS hypothesis using the single cointegration method.

TABLE 7.7: Cointegration Tests Results for Singapore (1970-2006)[105]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	I(1)***	I(0)**	Inconclusive

Single Equation Cointegration Approach[106]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	No	2	-0.08 [-0.66]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

105 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
106 t-values are given in squared-brackets.

Similar findings result when I use the multivariate cointegration test procedure. The Trace and the Maximum Eigenvalue test results, for both specifications of the Johansen ML cointegration test, produce a rank of zero. This contradicts the possibility of long-run cointegration between the two variables. Thus, once again, I reject the domestic BS model for Singapore.

7.2.8 Sri Lanka

The time plots for price and productivity series are plotted in FIGURE 7.1. The graphs are not very illuminating in the sense that nothing much is revealed from the periodic movements of two series. The two series sharply trend in opposite directions, till the year 1995. However, after this time, they co-move in a similar direction along with frequent intersections. This can be taken in support of their capability to establish cointegrating relationship in long-run.

Turning to the single equation cointegration tests in TABLE 7.8, for both the EG procedure and the ECM, the two tests fail to reject the null hypothesis of no cointegration at the 10 percent significance level. These findings suggest no long-run association between productivity differentials and the relative price of nontradables for Sri Lanka.

The multivariate cointegration model also could not detect any valid long-run association between the model variables. I obtain a rank zero for both the Trace and Maximum Eigenvalue tests, using both specifications of the Johansen ML cointegration approach. Thus, consistent with the single equation results, I find an absence of cointegration between relative prices of nontradables and its productivity-based determinant. I conclude there is no BS effect for Sri Lanka.

TABLE 7.8: Cointegration Tests Results for Sri Lanka (1981-2010)[107]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	I(1)***	I(1)***	I(1)
a	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[108]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	No	2	-0.03 [-0.74]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$pNT - pT$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

[107] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[108] t-values are given in squared-brackets.

7.2.9 Thailand

In FIGURE 7.1, the time plots for internal prices and productivity ratio visibly suggest lack of long-run association between two series. This is because (a) the series are trending in opposite direction, and (b) price series movements does not seem to be very promising in bringing adjustments to short-run misalignment for attaining long-run equilibrium.

My visual assessment is not supported by the single equation analysis in TABLE 7.9. For the EG residuals test, the associated tau-value when tested against the MacKinnon (1996) critical values suggests there there is a long-run relationship between productivity differentials and the relative price of nontradables for Thailand. The EC coefficient supports this finding, with an estimated EC coefficient of -0.17 (statistically significant), indicating that 17 percent of each period's fluctuations in the price ratio is corrected by self-movements, returning the series back to its equilibrium. However, the long-run BS effect has the wrong sign. To make things worse, both FMOLS and DOLS estimates of -0.30 are highly significant. This suggests that a rise in the country's relative productivity of tradables is associated with a depreciation in the relative price of nontradables, exactly the opposite of what we would expect from the BS effect.

In contrast, the Johansen ML cointegration tests for both Cases 3 and 4 do not find evidence that the series are cointegrated. Both of the model specifications indicate a rank of zero, demonstrating the lack of a long-run association between productivity differentials and relative price movements. Thus, as I have for all the previous country studies, I conclude that a domestic BS effect does not hold for Thailand.

TABLE 7.9: Cointegration Tests Results for Thailand (1971-2013)[109]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}$	Yes	Greater than I(1)	I(1)**	Inconclusive
a	Yes	I(1)***	I(1)**	I(1)

Single Equation Cointegration Approach[110]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC Coefficient	BS Coefficient		Does BS Effect Hold?
				FMOLS	DOLS	
$p^{NT} - p^{T}$	Yes	3	-0.17 [-2.40]	-0.30 [-10.88]	-0.30 [-9.67]	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}$	0	0	No

109 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
110 t-values are given in squared-brackets.

7.2.10 Summary of Individual Country Studies

The preceding sections of this chapter have dealt with examining the domestic version of the BS hypothesis for each of the subject Asian countries individually. A summary of the single equation and multivariate cointegration test results for the individual countries is reported in TABLE 7.10. The role of productivity gains in the tradable sector as a determinant of internal relative price movements of nontradables (internal real exchange rate) has found support in a number of non-Asian and Asian empirical studies on the BS hypothesis. While my findings stand in contrast to these studies, they are consistent with my earlier results about international productivity differentials and real exchange rates.

TABLE 7.10: Does the Domestic Version of Balassa-Samuelson Effect Hold? Summary of Results by Country

Country	Individual Results			Country Summary
	Single Equation Cointegration Method	*Multivariate Cointegration Method*		
		Case 3	*Case 4*	
Indonesia (1976-2013)	No	No	No	No
Japan (1970-2013)	No	No	No	No
Korea (1970-2013)	No	No	No	No
Malaysia (1980-2013)	No	No	No	No
Pakistan (1973-2008)	No	No	No	No
Philippines (1971-2013)	No	No	No	No
Singapore (1970-2006)	No	No	No	No
Sri Lanka (1981-2010)	No	No	No	No
Thailand (1971-2013)	No	No	No	No

7.2.11 Hong Kong

The country study of Hong Kong is discussed in a slightly different manner because – as noted previously -- Hong Kong does not follow a unique sectoral division. The empirical analysis uses four alternative types of sectoral disaggregations in order to cover the range of reasonable possibilities. Accordingly, the price and productivity measures are numbered from 1 to 4 to indicate the respective sectoral variants.

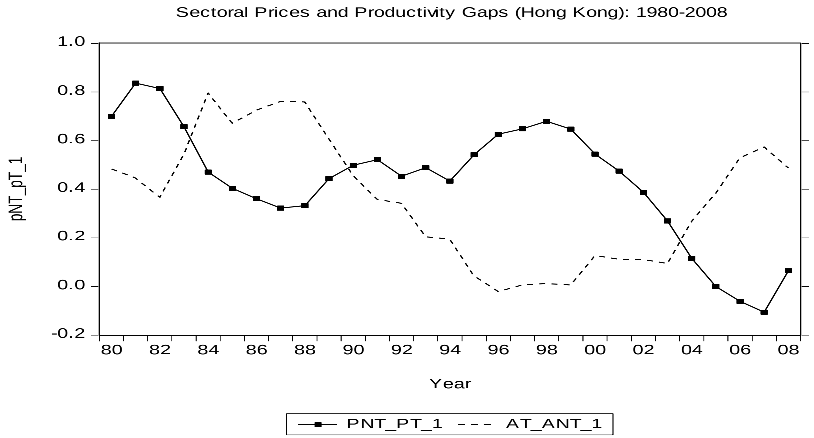
The time plots of model variables, for all four sectoral categories are given in FIGURE 7.2. The country does not seem to be an ideal candidate for establishing the BS effect. For three of the four sectoral divisions (Hong Kong_2 through Hong Kong_4), the two series are not co-moving in a common direction. In fact, the two series appear to be inversely related. Only for the first sectoral division, Hong Kong_1, do the two series behave in a manner consistent with the BS hypothesis. A more rigorous statistical analysis follows.

The empirical results are reported in TABLE 7.11. Starting with the single equation cointegration method, I find no evidence for the BS hypothesis using any of the four sectoral divisions. While classification 2 and 3 provide some evidence of cointegration between relative prices and productivities, the corresponding FMOLS and DOLS estimates of the BS coefficient are wrong-signed and, in all instances, highly significant. Thus, irrespective of the type of sectoral division, there is nothing here that supports the BS hypothesis.

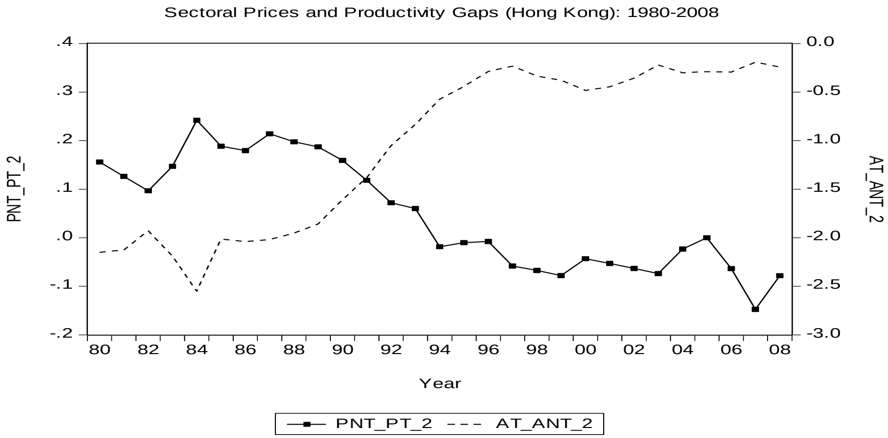
Similar results are obtained when the model is tested using the multivariate cointegration method. For all four sectoral divisions, a large majority of the test statistics (both Trace and Maximum Eigenvalue) concludes that the rank equals zero. The only exception occurs for the second sectoral division (cf. $p^{NT} - p^{T}\_2$). Here the Trace statistic finds evidence of a single cointegrating vector for Case 3 of the Johansen test. Accordingly, I run the respective VEC estimation to estimate the short- and long-run elements of the VECM.

Consistent with FIGURE 7.2, the VEC estimates indicate an inverse BS effect. The results meet the first pre-condition necessary for the existence of a valid BS effect. Deviations in the price series from long-run equilibrium are responded to with significant adjustments to return to equilibrium. But the second pre-condition, i.e., weak exogeneity, is not met, as can be seen from the second error correction coefficient (EC_a)in TABLE 7.11. Finally, the estimated coefficient on the BS variable is negative and highly significant, counter to the prediction of the BS hypothesis. Thus, the evidence across all sectoral divisions and using both single and multiple equation models points to the absence of a BS effect for Hong Kong.

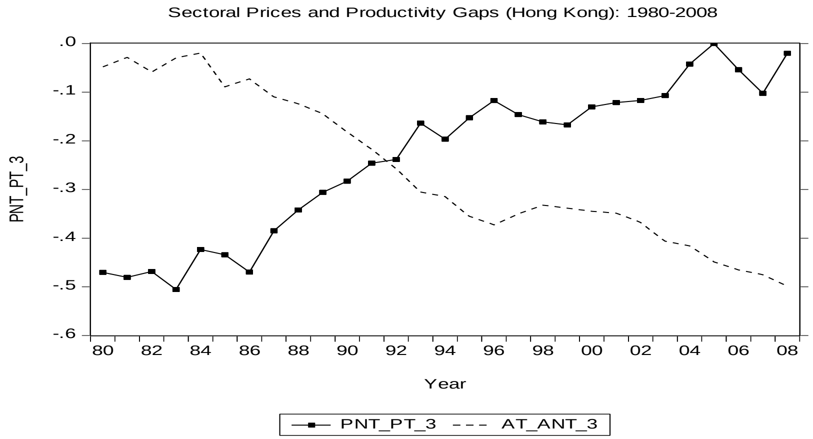
FIGURE 7.2: Plots for Domestic Sectoral Prices and Productivity Differentials (Hong Kong: 1978-2008)



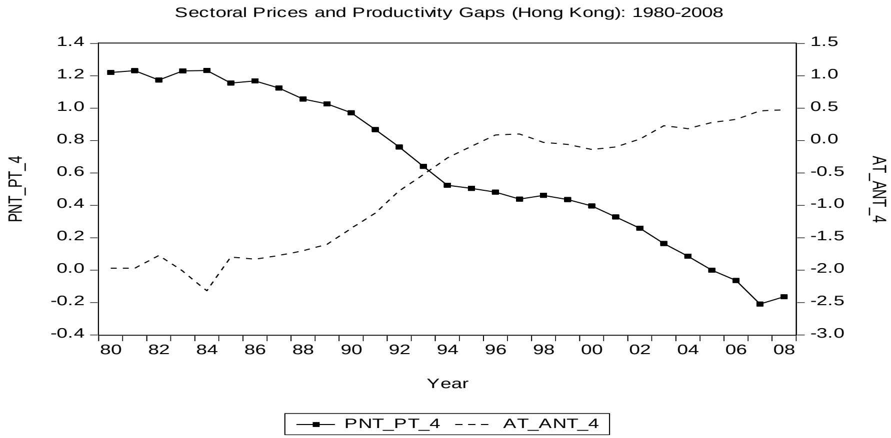
Hong Kong_1[111]



Hong Kong_2[112]



Hong Kong_3[113]



Hong Kong_4[114]

---

[111] Construction is the only nontradable sector here. Rest of all the sectors are treated as tradables.
[112] Mining, utilities, construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.
[113] Mining, utilities and construction are nontradable sectors whereas rest of all the sectors are treated as tradables.
[114] Construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.

TABLE 7.11: Cointegration Tests Results for Hong Kong (1978-2008)[115]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$p^{NT} - p^{T}\_1$	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
$p^{NT} - p^{T}\_2$	Yes	I(0)***	I(1)***	Inconclusive
$p^{NT} - p^{T}\_3$	Yes	I(1)***	I(1)***	I(1)
$p^{NT} - p^{T}\_4$	Yes	I(0)**	I(1)***	Inconclusive
a_1	Yes	I(1)***	I(1)***	I(1)
a_2	Yes	I(1)***	I(1)***	I(1)
a_3	Yes	I(1)***	I(1)***	I(1)
a_4	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[116]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of $p^{NT} - p^{T}$	EC[117] Coefficient	BS Coefficient FMOLS	BS Coefficient DOLS	Does BS Effect Hold?
$p^{NT} - p^{T}\_1$	Yes	2	-0.15 [-1.54]	-	-	No
$p^{NT} - p^{T}\_2$	No	1	-1.24 [-5.58]	-0.13 [-14.00]	-0.14 [-14.62]	No
$p^{NT} - p^{T}\_3$	Yes	1	-0.81 [-2.09]	-0.20 [-24.43]	-0.21 [-28.26]	No
$p^{NT} - p^{T}\_4$	No	1	0.03 [0.32]	-	-	No

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

Version of Price Ratio	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}\_1$	0	0	No
$p^{NT} - p^{T}\_2$	1	0	See below
$p^{NT} - p^{T}\_3$	0	0	No
$p^{NT} - p^{T}\_4$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

Version of Price Ratio	Trace Statistics	Max Eigen Value	Does BS Effect Hold?
$p^{NT} - p^{T}\_1$	0	0	No
$p^{NT} - p^{T}\_2$	0	0	No
$p^{NT} - p^{T}\_3$	0	0	No
$p^{NT} - p^{T}\_4$	0	0	No

[115] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[116] t-values are given in squared-brackets.
[117] The four test regressions also contain first and second lagged-difference of a.

Vector Error Correction Model[118]

Version of Price Ratio	Lags	White Noise Residuals	$EC_{p^{NT}-p^{T}}$	EC_a	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
$p^{NT} - p^{T}$_2	0	Yes	-0.67 [-3.54]	2.27 [2.70]	-0.13 [-10.36]	No

[118] t-values are given in squared-brackets.

7.3 Panel Data Estimation Results

Panel unit root test results for the Fisher-ADF, Fisher-PP and Hadri tests for the relative price and productivity series are reported in the top panel of TABLE 7.12. For relative sectoral prices, all three tests indicate the series to be a unit root process with high statistical significance. The results for the productivity series are confusing, as the three tests all produce different results. As a result, I cannot confidently conclude the actual order of integration for this series.

The second and third panels of TABLE 7.12 report test results for the Pedroni residual based cointegration test and Johansen-Fisher panel cointegration model. Discussing the test statistics obtained from the Pedroni cointegration test first, I opted for automatic lag selection. All seven test statistics fail to reject the null hypothesis of no cointegration between model variables. Thus, no support could be obtained from the Pedroni cointegration test in favour of a domestic BS effect for full panel of nine countries.

Turning now to the results obtained from the Fisher-Johansen cointegration tests, the test requires users to specify lag length. I select lag length through Panel VAR, using the minimum SIC value as my criterion. The Trace and Maximum Eigenvalue values of the two specifications are the same within case, but different across cases. Specification 3 of the test produces a rank of 2, indicating that the variables are stationary in levels. Specification 4 yields a rank of zero, indicating that the variables do not share a common long-run relationship. Thus, I conclude that the ML based rank cointegration tests provide no support for the domestic version of the BS hypothesis.

In conclusion my panel results are consistent with my individual country results. Overwhelming, I find robust and consistent evidence against the domestic version of the BS hypothesis.

TABLE 7.12: Summary of Results for Panel Unit Root and Cointegration Tests for Domestic Balassa-Samuelson Effect[119,120]

Panel Unit Root Test Results (Order of Integration as Determined by)

Variables	Fisher-ADF	Fisher-PP	Hadri	Conclusion
$p^{NT} - p^{T}$	I (1)***	I (1)***	I (1)***	I (1)
a	I (0)**	I (1)***	Greater than I (1)	Inconclusive

Pedroni Panel Cointegration Test Results[121]

Common AR Coefficients (Within Dimension)				*Individual AR Coefficients (Between Dimension)*			
Panel v Statistics	Panel ρ Statistics	Panel PP Statistics	Panel ADF Statistics	Group ρ Statistics	Group PP Statistics	Group ADF Statistics	Does BS Effect Hold?
-1.37	1.84	1.66	1.19	1.70	0.99	0.91	No

Johansen-Fisher Panel Cointegration Test Results[122,123]

Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigen Stat)	Does BS Effect Hold?
Case 3: Intercept (no trend) in cointegrating equation and VAR		
2	2	No
Case 4: Intercept and trend in cointegrating equation-no trend in VAR		
0	0	No

[119] ***, ** and * are representing significance of sample statistics at 1%, 5% and 10% levels respectively.
[120] Hong Kong is omitted from panel estimations.
[121] Pedroni panel cointegration is a test for null of no cointegration in both homogenous and heterogeneous panels. The test statistics are standardized and asymptotically normally distributed. See Pedroni (1995, 1999) for further details.
[122] The test is maximum likelihood based rank test.
[123] Lag selection is done through SIC under panel VAR.

7.3.1 Summary of Panel Data Results

The conclusions drawn from pooled data estimators are no different from those obtained from individual country analysis. The probable existence of the domestic version of BS model can be declined with high level of certainty, as evident from two panel data estimators' results. Thus, even after conducting a rigorous empirical examination, using alternative cointegration tests of time-series and panel data estimations methods, absolute lack of support is found in favour of domestic BS hypothesis. In the forthcoming chapters of the dissertation, I will resume my analysis of standard (international) version of the model, in its modified version, by redefining the model for its over-restrictive and highly idealist theoretical assumptions.

TABLE 7.13: Does Domestic Version of Balassa-Samuelson Effect Hold? Summary of Results for Panel Cointegration Tests

Test of Cointegration			
Pedroni Residual Based Panel Cointegration Test	Johansen-Fisher Panel Cointegration Test		Conclusion
	Case 3	Case 4	
No	No	No	No

APPENDIX-C

EViews Programming Code for Indonesia

```
wfopen "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-7\Indonesia.wf1"

'***************************************************************************
'Group Plot for pNTpT, RER_DEF, RER_DEF_NT and aTaNT
'***************************************************************************
group gA PNT_PT AT_ANT
freeze(group_plot) gA.line(x)
group_plot.setelem(1) lcolor(black) symbol(7) lpat(1)
group_plot.setelem(2) lcolor(black) symbol(4) lpat(1)
group_plot.setelem(3) lcolor(black) symbol(1) lpat(1)
group_plot.setelem(3) lcolor(black)
group_plot.options linepat
group_plot.addtext(t) Sectoral Prices and Productivity Gap (Indonesia & U.S): 1976-2013
group_plot.addtext(b) Year
group_plot.addtext(l) PNT_PT
group_plot.addtext(r) AT_ANT

'************************************************************
'************************************************************
create y 1976 2013
'importing data from Excel for Indonesia
import "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-7\Chapter 7.xlsx" range="Indonesia"

'**************************************************************************************************
'CASE-1: ESTIMATING BALASSA-SAMUELSON EFFECT FOR pNTpT & aTaNT
'**************************************************************************************************
'*************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'*************************************************************
'*********************************
'Graph for Indonesia's pNTpT
'*********************************

genr pNTpT = pNTpT
freeze(figure_pNTpT) pNTpT.line
figure_pNTpT.addtext(t) pNTpT (Indonesia):  1976-2013
figure_pNTpT.addtext(b) Year
figure_pNTpT.addtext(l) pNTpT
figure_pNTpT.legend(off)

'We see from the FIGURE that pNTpT has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'************************************************
'ADF Unit Root Test for Indonesia's pNTpT
'************************************************

freeze(table_7_1_1_pNTpT_adf) pNTpT.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 2.  The unit root test produces a t-value of -1.95 which is greater than our 5% criterion -3.54.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
```

```
freeze(mode=overwrite,pNTpT_adf) pNTpT.uroot(adf,const,trend,info=sic)
freeze(pNTpT_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the pNTpT series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr pNTpTdiff = d(pNTpT)
freeze(figure_pNTpTdiff) pNTpTdiff.line
figure_pNTpTdiff.addtext(t) dpNTpT (Indonesia):  1976-2013
figure_pNTpTdiff.addtext(b) Year
figure_pNTpTdiff.addtext(l) DpNTpT
figure_pNTpTdiff.legend(off)
```

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

```
genr pNTpTdiff = d(pNTpT)
freeze(table_7_1_2_pNTpTdiff1_adf) pNTpTdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -4.91 which is now smaller than our 5% criterion -2.94. Thus, we may now reject the null of non-stationarity in first differenced series of pNTpT. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,pNTpTdiff1_adf) pNTpTdiff.uroot(adf,const,info=sic)
freeze(pNTpTdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the pNTpT series is I(1).

'***
'DF-GLS Unit Root Test for Indonesia's pNTpT
'***

```
freeze(table_7_1_3_pNTpT_dfgls) pNTpT.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -1.45 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of unit root.

'Now let's see if the series is difference stationary or not

```
genr pNTpTdiff = d(pNTpT)
freeze(table_7_1_4_pNTpTdiff1_dfgls_d) pNTpTdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -4.57 which is now smaller than our 5% criterion -1.95. Thus, we may reject the null of non-stationarity in first differenced series of pNTpT.

''Putting it all together, I conclude that the pNTpT series is I(1), a finding compatible with my ADF test results.

'***
'Graph for Indonesia's Productivity (aTaNT)
'***

```
genr aTaNT = aTaNT
freeze(figureaTaNT) aTaNT.line
figureaTaNT.addtext(t) aTaNT (Indonesia):  1976-2013
figureaTaNT.addtext(b) Year
figureaTaNT.addtext(l) aTaNT
figureaTaNT.legend(off)
```

```
'We see from the FIGURE that aTaNT has time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'*********************************************************
'ADF Unit Root Test for Indonesia's Productivity
'*********************************************************

freeze(table_7_1_1_aTaNT_adf) aTaNT.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p =0.  The unit root test produces a t-value of -5.24 which is smaller than our 5% criterion -3.53.  Thus, at this point, we can reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,aTaNT_adf) aTaNT.uroot(adf,const,trend,info=sic)
freeze(aTaNT_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the aTaNT series is level stationary.

'**************************************************************
'DF-GLS Unit Root Test for Indonesia's Productivity
'**************************************************************

freeze(table_7_1_3_aTaNT_dfgls) aTaNT.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 1.  The unit root test produces a t-value of -1.35 which is greater than our 5% criterion -3.19.  Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr aTaNTdiff = d(aTaNT)
freeze(table_7_1_4_aTaNT1diff1_dfgls) aTaNTdiff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 1.  The unit root test produces a t-value of -0.91 which is still greater than our 5% criterion -1.95. Thus, we may not reject the null of non-stationarity in first differenced series of aTaNT.

''Putting it all together, I conclude that the aTaNT series is greater than I(1), a finding incompatible with my ADF test results.

'**********************************************
'Single Equation Cointegration Methods
'**********************************************

'************************************************************
''Graph the suspected cointegrated series together
'************************************************************

'The first step is to plot a graph of the suspected series.  This is very important!

group g1 pNTpT aTaNT
freeze(figure7_1a) g1.line(x)
figure7_1a.setelem(1) lcolor(black)
figure7_1a.setelem(2) lcolor(black) lpat(8)
figure7_1a.options linepat
figure7_1a.addtext(t) pNTpT and aTaNT (Indonesia & U.S): 1976-2013
figure7_1a.addtext(b) Year
figure7_1a.addtext(l) pNTpT
figure7_1a.addtext(r) aTaNT
''*********************************************************
''S1.A.Engle-Granger Approach to Cointegration
'*********************************************************

freeze(table_7_1_egc) g1.coint(method=eg)
```

```
'The null hypothesis will not be rejected as suggested by sample statistics.

''**********************************************
''S1.B.Error Correction Model (ECM)
'***********************************************
''**************************************************
'Selecting the number of lags in the VAR
'***************************************************

'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var table_7_1_var1.ls 1 4   g1
freeze(table_7_1_var1_lagtest1) table_7_1_var1.laglen(4)
freeze(table_7_1_var1_lagtest2) table_7_1_var1.testlags

'The laglength test above indicates that the VAR has 1 lag. But the residuals were not white noise at 1 lag. So, I had to raise the number of lags from 1 to 3 to obtain white residuals.

var table_7_1_var2.ls 1 3  g1
freeze(table_7_1_var2_artest1) table_7_1_var2.correl
freeze(table_7_1_var2_artest2) table_7_1_var2.qstats(12)
freeze(table_7_1_var2_artest3) table_7_1_var2.arlm(12)

'The residuals are absolutely white noise.

'We now try different lags of d(aTaNT), comparing SIC values across specifications.

genr resid = 0
equation eg.ls pNTpT c aTaNT
genr ec = resid

var table_7_1_eg2a.ls 0 0 d(pNTpT)   @  c ec(-1) d(pNTpT(-1)) d(pNTpT(-2)) d(pNTpT(-3))

var table_7_1_eg2b.ls 0 0 d(pNTpT)   @  c ec(-1) d(pNTpT(-1)) d(pNTpT(-2)) d(pNTpT(-3)) d(aTaNT(-1))

var table_7_1_eg2c.ls 0 0 d(pNTpT)   @  c ec(-1) d(pNTpT(-1)) d(pNTpT(-2)) d(pNTpT(-3)) d(aTaNT(-1)) d(aTaNT(-2))

'The evidence suggests that Model C is best.  Now we test that model for serial correlation.

var table_7_1_eg2c.ls 0 0 d(pNTpT)   @   c ec(-1) d(pNTpT(-1)) d(pNTpT(-2)) d(pNTpT(-3)) d(aTaNT(-1)) d(aTaNT(-2))
freeze(table_7_1_eg2c1_artest1) table_7_1_eg2c.correl
freeze(table_7_1_eg2c2_artest2) table_7_1_eg2c.qstats(12)
freeze(table_7_1_eg2c3_artest3) table_7_1_eg2c.arlm(12)

'The residuals are absolutely white noise.

''**************************
''Estimating EC Model
'***************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_7_1_ecm.ls(n) d(pNTpT) c ec(-1) d(pNTpT(-1)) d(pNTpT(-2)) d(pNTpT(-3)) d(aTaNT(-1)) d(aTaNT(-2))

'Note that the SR effect is significant as the error correction coefficient -0.09 is statistically significant at 10% significance level.

''**************************************************
''S2.A & S2.B: Obtaining LR Coefficients
'***************************************************
```

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Indonesia against U.S.

equation table_7_1_LReqn_fmols.cointreg(method=fmols) pNTpT aTaNT

equation table_7_1_LReqn_dols.cointreg(method=dols, trend=constant, lag=2,lead=2) pNTpT aTaNT

'The BS coefficient obtained through FMOLS and DOLS estimators are 0.15 and -0.04, i.e., the long-run BS coefficients are bearing undired signs and/or are statistically insignificant. Thus, there is 'NO' evidence in support of BS effect existing for Indonesia.

''**
'Multivariate Cointegration Approach
'**

''**
''Check if the VAR (2) model is dynamically stable
'**
freeze(table_7_1_var2_varstable) table_7_1_var2.arroots(graph)

'The model is dynamically stable.

''**
''M1.A & M1.B: Identifying the number of cointegrating vectors
'**

'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_7_1_var2_coint) table_7_1_var2.coint(s,3)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. All the results indicate 0 cointegrating vectors.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

''***
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'***

' For estimating the LR relationship, corresponding VEC command is:

var table_7_1_vecd.ec(d,1) 1 2 pNTpT aTaNT

'CONCLUSION: I conclude that pNTpT and aTaNT are not cointegrated in the Indonesia's data.

CHAPTER 8: ESTIMATING THE MODIFIED VERSION OF THE BALASSA-SAMUELSON HYPOTHESIS

8.1 Motivation

The preceding chapters found little empirical evidence in support of the Balassa-Samuelson (BS) effect for the subject Asian economies. The associated findings were robust over two alternative sectoral classification schemes, three distinct measures of real exchange rates, different empirical estimation methods and two alternative theoretical specifications of the hypothesis. Each time, the model results remained unchanged, demonstrating the absence of a valid long-run association between inter-country relative gains in tradables productivity and real exchange rate movements, contrary to the theoretical predictions of the BS hypothesis.

The difference between my findings and other studies that support the BS hypothesis motivates me to address the hypothesis from a new perspective. The BS hypothesis is often criticized for its highly idealistic assumptions. A number of studies test the model by relaxing its over-simplified and unrealistic assumptions, and obtain different results from what they would have obtained otherwise. For Asia-Pacific Economic Cooperation (APEC) countries, Ito et al. (1999) discovered that real exchange rate movements in the past three decades have largely been determined by inter-country tradables' price movements. Such a revelation casts a shadow over one of the fundamental assumptions underlying the BS hypothesis. Namely, that Purchasing Power Parity (PPP) holds for tradables prices across countries.

The existence of PPP for inter-country tradables prices has also been empirically challenged by MacDonald and Ricci (2001) for European countries, and Thomas and King (2008) for East Asian developing states. In fact, a wide range of literature documents that inter-country tradables prices sizeably and persistently deviate from long-run PPP (Canzoneri et al., 1999; Egert, 2002b; Egert et al., 2003; Kovacs, 2003; Lojschova, 2003; Blaszkiewicz et al., 2004; MacDonald and Ricci, 2005; Lee and Tang, 2007; Garcia-Solanes et al., 2008).

Schmillen (2013) empirically evaluates another fundamental assumption of the BS hypothesis for OECD and Central and Eastern European economies; i.e., homogenous labour markets and perfect wage equalization across sectors. His results indicate that the assumption of wage equalization across sectors of production does not hold. This lends renewed support to the view that multi-sector open-economy macroeconomic models, especially those concerned with the BS hypothesis, might benefit from weakening the assumption of homogeneous labour markets. The violation of the assumption of inter-industry wage equalization also finds support from other studies (Strauss and Ferris, 1996; Strauss, 1997, 1998; Nenovsky and Dimitrova, 2002; Lee, 2005).

MacDonald and Ricci (2005) attribute real exchange rate departures from long-run equilibrium to the consumption of nontradables in the production process of tradables. They examine the role of distribution sector, typically categorized as a nontradable industry. They identify the co-existence of the distribution sector with the usual productivity-driven BS effect (relative bias in the domestic productivity of tradables) as a significant determinant of real exchange rate depreciation in ten OECD countries. Estrada and Lopez-Salido (2004) link the problem of dual inflation in Spain with rising mark-ups for services and manufacturing. Contrary to one of the assumptions of the BS hypothesis, i.e., that the capital-labour ratio will remain constant across sectors, the authors find the evolution of mark-ups in the services and manufacturing sectors to be the main driver of Spanish inflation vis-a-vis Europe.

Based on the preceding studies, I re-examine the BS hypothesis using a modified version of the model. I relax the standard (international) version of the BS hypothesis, allowing for the failure of the law of one price (LOP) in tradables. The new estimable equation for testing the BS effect will be set up in a way to allow home and U.S. tradables prices to deviate from the LOP. Such an approach will not only provide me with an opportunity to obtain more reliable estimates of the BS effect, but will also provide an insight into the role of tradables prices in determining real exchange rate trend departures from long-run equilibrium.

8.2 Law of One Price and Purchasing Power Parity in Tradables

The LOP hypothesizes that identical goods, when their national prices are measured in a common currency, should be sold for a common price when traded at different geographical locations. Generalizing the phenomenon by aggregating across various intra-sector and inter-sector tradable goods yields the notion of Purchasing Power Parity (PPP) in tradables; a common basket of goods consisting of tradables (only), exchangeable internationally for a common price.

The LOP is often regarded as a long-run phenomenon for arbitrage in international goods markets. The idea explains that arbitrage opportunities that may be observed in international goods markets but are of temporary and short-lived nature, offset by the adjustments induced by either nominal exchange rate and/or price movements, thus allowing LOP to re-establish in long-run.

8.2.1 Tradables' Purchasing Power Parity and Balassa-Samuelson Effect

One of the vital assumptions of the standard theoretical model underlying the BS hypothesis is that Purchasing Power Parity (PPP) holds for the inter-country traded sectors. The traditional BS effect occurs when domestic and foreign tradables are perfect substitutes and PPP holds for tradable goods, i.e., ($rer^T = 1$ or some other constant value). This implies that any deviations of rer^T from one will be transient and temporary. PPP holding between home and foreign tradables, a productivity improvement in the domestic tradable sector will drive up the wages for the entire economy and thus the labour cost for the nontradable sector as well. The relative price of the nontradable sector at home will rise consequently, leading to trend appreciation in the country's real exchange rate.

I begin by revising the model to relax the assumption of PPP. To do that, I revisit the accounting framework for the real exchange rate based on the theoretical model of Chapter Two.

Using equation 2.13 of Chapter 2, the real exchange rate can be written in logarithmic form as:

$$(8.1) \quad rer = (e + p^{T*} - p^T) - \beta\{(p^{NT} - p^T) - (p^{NT*} - p^{T*})\}.$$

The first term on the right hand side of Equation (8.1) controls for influences of the relative prices of tradables on the real exchange rate. If PPP holds between tradables prices of two countries, this term equals zero. If PPP does not hold, inter-country biased relative productivity of tradables can cause trend movements in real exchange rates through both nontraded and traded sector price movements. This might undermine the BS effect which is presumed to work solely through nontraded sector prices.

Equation (2.14-a) of Chapter Two transforms the above equation to model real exchange rates against the traded sector price differential and the relative sectoral productivity differential between home and U.S. as follows:

$$(8.2) \quad rer = (e + p^{T*} - p^T) - \beta\left\{\left({}^{\gamma}/_{\delta}\, a^T - a^{NT}\right) - \left({}^{\gamma}/_{\delta}\, a^{T*} - a^{NT*}\right)\right\}$$

where $rer = e + p^* - p$. e is the bilateral nominal exchange rate between the home and foreign country. p and p^* typically refer to prices of a basket of nontradable goods (or at minimum a basket largely comprised of nontradable items), whose prices are determined through domestic market forces. For both the home and foreign markets, p and p^* are measured through nontraded sector, value-added deflators in my analysis. $(e + p^{T*} - p^T)$ is analogous to traded sector prices based on the real exchange rate (rer_def_t). Hence, equation (8.2) can be re-written as:

(8.3) $\quad rer\_def\_nt = (rer\_def\_t) - \beta\{(a^T - a^{NT}) - (a^{T*} - a^{NT*})\}$[124]

Equation (8.3) establishes the long-run relationship between inter-country biased sectoral productivity and real exchange rate (NT) movements, besides allowing for traded sector prices to deviate from PPP equilibrium. This leads to the estimable version of the model:

(8.4) $$rer\_def\_nt = \alpha + \vartheta(rer\_def\_t) - \beta\tilde{a} + \mu$$

Equation (8.4) is the final estimable equation for the modified version of the BS effect, allowing deviations from PPP in tradables prices in the home and U.S. markets. This implies that the change in the rer_def_nt in an accession country depends on inter-country sectoral productivity gap ($\tilde{a}$) as well as tradables' price differentials (rer_def_t) between two countries. This is in contrast with traditional BS effect where rer_def_nt is a function of $\tilde{a}$ only. ϑ is responsible for capturing the deviations of relative tradables price from long-run PPP, thus explaining rer_def_nt trend movements. A long-run coefficient value of ϑ, sizeably different from one (with appropriate statistical significance), will be evident of violating the condition of PPP for inter-country tradables[125]. However, this will not invalidate the BS effect (if $\tilde{a}$ holds a positive a valid long-run coefficient)[126]. As always, β is the long-run BS slope coefficient, expected to bear a positive and

[124] Labor intensities are assumed to be constant across sectors, implying $\gamma/_{\delta} = 1$

[125] Since rer_def_t is not the main variable of interest, its statistics are always reported in black, irrespective of their statistical significance (insignificance).

[126] For individual country analysis, the condition of weak exogeneity of model regressors, under multivariate cointegration approach (please refer to Chapter-5, TABLE 5.2, M2.B.) for establishing valid BS effect is no more effective.

statistically significant value for establishing a valid BS effect. μ represents the white noise model residuals.

8.2.2 Empirical Evidence on the Assumption of Tradables' PPP

My previous inability to find empirical evidence in favour of a valid BS effect for Asian countries may have been due to the built-in assumption of PPP for tradables. It is important to note that if this assumption fails to hold, it would not challenge the proposed BS mechanism. It merely causes the conventional model to be under-specified (Thomas and King, 2008). Under the BS hypothesis, trend departures of tradables prices from PPP do not hinder the ability of the inter-country productivity gap to generate trend deviations in real exchange rate from its long-run equilibrium. However, failure of tradables PPP will affect rer^T, and failure to control for this effect could cause deficient estimates of the BS effect.

I now turn to the empirical evidence on the validity of the LOP and/or PPP for tradables. A number of studies, exploring LOP/PPP for tradables, find that tradables prices are highly volatile with persistent and long-lived deviations (Isard, 1977; Richardson, 1978; Knetter 1989, 1993). Market structure, product differentiation, trade restrictions, distance between markets, transportation costs, nominal exchange rates, etc. are some of the factors empirically tested for preventing relative tradables prices to converge and thus establish LOP/PPP in the long-run.

Isard (1977) confirms the invalidity of the LOP for Canada, Germany, Japan and U.S. using low level to highly disaggregated data. Using a straightforward OLS regression model, the study finds that deviations from the LOP are substantial and persistent. Relative tradables' prices are observed to be highly sensitive to nominal exchange rate movements for all the accession countries (except Canada). Similar to Isard (1977), Richardson (1978) does not find support for the LOP while testing the phenomenon for Canada and the U.S. Employing high quality disaggregated data for a set of 22 commodities, the author rejects the equi-proportionate relationship between traded sector prices. Canadian prices are found to be highly sensitive to nominal exchange rate movements. Engel and Rogers (1996) attribute much of the variation in the CPI based real exchange rate of Canada to the non-existence of the LOP. Testing monthly CPI data for U.S. and Canadian cities for 14 commodities over the period 1978 to 1993, the authors find that physical distance and border play a significant role in explaining the failure of the LOP.

Sarno et al. (2004) conduct an empirical investigation in support of non-zero international transaction costs, explaining trend deviations of tradables' prices from the LOP. Constructing real exchange rates using quarterly price deflators for nine sectors (from 1974 to 1993), the analysis is done for France, Germany, Italy, Japan, UK and the U.S. (as the bench mark country). Using the Threshold Autoregressive (TAR) model, the authors find that deviations from the LOP dissipate in a non-linear fashion. The deviations from the LOP are found to be mean reverting with a credible speed of convergence. This suggests that deviations from LOP are not long-lived or persistent.

Only a handful of studies have tested the assumption of tradables' PPP in the context of the BS hypothesis for Asia. Chapter Four includes a separate section, containing a detailed discussion of these studies, highlighting their important findings. The existing empirical evidence on this issue is too limited to develop a clear understating on the sensitivity of Asian real exchange rates towards trend departures of tradables prices from PPP equilibrium. In upcoming sections of the chapter, I extend research in this area by relaxing the assumption of tradables' PPP. I continue to test the BS hypothesis, but use a modified model that does not assume PPP between home (Asia) and U.S. traded sector prices. Consistent with earlier chapters, this modified version of the hypothesis will be tested for individual countries before pooling the data as a panel.

8.3 Country Studies

Analogous to preceding chapters, this section will discuss country by country results of the modified version of the BS hypothesis, incorporating the possible absence of PPP between inter-country tradables prices. The modified version of the model is tested empirically using equation (8.4) given in Section 8.3. The results are obtained using the same general procedures as in preceding chapters. The EViews program for Korea is attached in the Appendix to this chapter to provide an example of the programming code used to derive the results reported in this chapter.

Before modelling the role of inter-country traded sector prices in BS hypothesis, I verify the existence (inexistence) of PPP between traded sector prices of home and U.S. The test of PPP is conducted for each country individually. Those countries are then tested for the modified version of BS hypothesis (only) for which, valid statistical support is yielded in favour of inexistence of PPP for their tradables. The following section lays out a detailed discussion on the verification of the PPP assumption for Asia and the U.S., consisting of a discussion of testing and estimation procedures, and subsequent results.

8.3.1 Testing the Assumption of PPP for Inter-country Tradables Prices

The idea of PPP originally stems from the Law of One Price (LOP). The LOP hypothesizes that identical goods, when their national prices are measured in a common currency, should be sold for a common price, while being traded at different geographical locations. Generalizing the phenomenon by aggregating across various intra-sector and inter-sector tradables yields the notion of Purchasing Power Parity (PPP) in tradables; a common basket of goods consisting of tradables (only), traded across borders for a similar price.

Theoretically, for home and international tradables, with dissimilar currencies and different locations of production, and ignoring transaction costs, the law of PPP can be stated as:

(8.5-a) $p^T = e + p^{T*}$,

where p^T and p^{T*} are the aggregate prices of home (Asia) and international (U.S.) tradables, expressed in their national currencies. The operational (estimable) version of the model is given as:

(8.5-b) $p^T = \gamma + \theta(e + p^{T*}) + \varepsilon$

The validity of the assumption of PPP for tradables is determined through testing equation (8.5-b). The equation is tested for two distinct but inter-linked conditions, stated below.

S1. In the long-run, home prices of tradables should co-move with their foreign counterpart.

The first condition serves as a necessary condition for validating the assumption of PPP. It implies that long-run co-movement amongst inter-country tradables prices can be taken as evidence in support of PPP. Home country (Asia) being a small open economy, its traded sector prices may retain their individuality in the short-run. But in the long-run, it does not have enough market power to influence foreign prices (U.S.) and hence, will follow the trend movements of U.S. prices. Thus, in the long-run, the tradables prices of the home country are largely determined by foreign prices, thus satisfying the law of tradables PPP.

The valid long-run co-movement in home and U.S. traded sector prices will be verified using two cointegration regression estimators; FMOLS and DOLS. If both estimators produce a statistically significant long-run relationship between the two model variables (p^T and $e + p^{T*}$), I will take this as strong evidence in support of long-run co-movement ("YES). If only one of the two estimators supports cointegration, I will take this as mixed evidence in support of long-run co-movement ("MIXED"). In either case (YES or MIXED), I will proceed with testing the second model condition (S2). On the other hand, if both cointegration estimators produce a statistically insignificant long-run relationship, I will conclude that the two price variables do not have a long-run relationship ("NO"), and I will not proceed with testing the second condition (given below).

S2. Home prices bear a direct and equi-proportionate relationship with their foreign counter-part.

This second condition will be tested formally through a Wald coefficient test. The null and alternative hypotheses are given as follows:

$$Null\ Hypothesis: H_0: \hat{\theta} = 1$$

$$Alternative\ Hypothesis: H_1: \hat{\theta} \neq 1$$

Strong evidence in favour of PPP exists if both cointegration estimation procedures fail to reject the null hypothesis above (“YES”). In this case, I will conclude that the modified version of the BS model is not “permissible.” If both of the cointegration estimation procedures reject the null hypothesis, I will take this as strong evidence against PPP (“NO”) and conclude that estimation of the modified BS model is permissible. And if one of the cointegration estimation procedures rejects the null hypothesis, while the other fails to reject it (“MIXED”), I will also conclude that this is sufficient support to warrant estimation of the modified BS model.

TABLE 8.1 summarizes the conditions above necessary to validate (invalidate) the assumption of PPP for tradables, in the individual country studies. The last column of the table shows if the estimation of modified version of BS hypothesis is permissible or not, under different combinations of S1 and S2.

TABLE 8.1: Conditions[127] for Verifying PPP between Inter-Country Tradables Prices

S1-Establishing Cointegration (FMOLS & DOLS) $p^T = \gamma + \theta(e + p^{T*}) + \varepsilon$ H_1: *Cointegration exists* θ *is statistically significant*	**S2-Testing for Equi-proportionate Relationship (Wald Coefficient Test)** $H_0: \hat{\theta} = 1$ H_0: Equi-proportionate relationship between p^T and $e + p^{T*}$	**Is Estimation of Modified BS Model Permissible?**
YES - Absolute support for H_1	'YES' absolute support for H_0	No
	'MIXED' support for H_0	Yes
	'NO' support for H_0	Yes
Mixed support for H_1	'YES' absolute support for H_0	No
	'MIXED' support for H_0	Yes
	'NO' support for H_0	Yes
No support for H_1	NA	Yes

The country estimates for equation (8.5-b) are reported in TABLE 8.2. As a prerequisite, the two model time series (p^T and $e + p^{T*}$) are tested for their order of integration, using ADF and DF-GLS unit root tests. For each country, the two model conditions (S1 and S2) are tested individually. FMOLS and DOLS estimates with subsequent F-statistics, obtained through a Wald coefficient test, are reported below the unit root test results. A 10-percent significance level will be used in testing hypotheses associated with S1 and S2.

[127] In Section 8.6 of the chapter, while conducting pooled data analysis, the existence (inexistence) of PPP will be confirmed against same two conditions. S1 and S2 will be titled as P1 and P2, respectively.

TABLE 8.2: Results for Testing the Assumption of PPP for Inter-Country Tradables Prices[128,129,130]

Country	Verification of Model Conditions				Does PPP Hold for Tradables Prices?
Indonesia	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	Yes
		p^T	I(1)	I(1)	
		$e + p^{T*}$	I(1)	I(1)	
	S1		**FMOLS**	**DOLS**	
			1.03 [16.62]	1.00 [15.01]	
	S2		**F-Statistics from Wald Test**		
			0.21 (0.65)	0.00 (0.99)	
Japan	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	No
		p^T	Greater than I(1)	I(1)	
		$e + p^{T*}$	I(1)	I(0)	
	S1		**FMOLS**	**DOLS**	
			0.04 [0.20]	0.07 [0.30]	
	S2		**F-Statistics from Wald Test**		
			-	-	
Korea	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	No
		p^T	I(0)	I(1)	
		$e + p^{T*}$	I(1)	I(1)	
	S1		**FMOLS**	**DOLS**	
			1.38 [24.95]	1.33 [25.30]	
	S2		**F-Statistics from Wald Test**		
			47.84 (0.00)	28.03 (0.00)	

[128] t-values are given in squared-brackets.
[129] p-values are given in parenthesis.
[130] Sample statistics of two unit root tests are tested against the null hypothesis at 5 percent significance level.

Country	**Verification of Model Conditions**				**Does PPP Hold for Tradables Prices?**
Malaysia	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	Yes
		p^T	I(0)	I(1)	
		$e + p^{T*}$	I(1)	I(1)	
	S1	**FMOLS**		**DOLS**	
		1.01 [5.84]		1.02 [5.58]	
	S2	**F-Statistics from Wald Test**			
		0.03 (0.97)		0.10 (0.92)	
Pakistan	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	No
		p^T	Greater than I(1)	I(0)	
		$e + p^{T*}$	I(1)	I(1)	
	S1	**FMOLS**		**DOLS**	
		0.89 [23.13]		0.86 [24.69]	
	S2	**F-Statistics from Wald Test**			
		7.33 (0.01)		15.12 (0.00)	
Philippines	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	No
		p^T	I(1)	I(1)	
		$e + p^{T*}$	I(1)	I(1)	
	S1	**FMOLS**		**DOLS**	
		1.10 [34.35]		1.06 [34.38]	
	S2	**F-Statistics from Wald Test**			
		3.03 (0.00)		2.10 (0.04)	
Singapore	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	Mixed
		p^T	I(1)	I(1)	
		$e + p^{T*}$	I(1)	I(1)	
	S1	**FMOLS**		**DOLS**	
		1.90 [4.43]		1.72 [3.81]	
	S2	**F-Statistics from Wald Test**			
		2.10 (0.04)		1.59 (0.12)	

<table>
<tr><th>Country</th><th colspan="4">Verification of Model Conditions</th><th>Does PPP Hold for Tradables Prices?</th></tr>
<tr><td rowspan="9">Sri Lanka</td><td>Unit Root Testing</td><td>Variables</td><td>ADF</td><td>DF-GLS</td><td rowspan="9">Mixed</td></tr>
<tr><td></td><td>p^T</td><td>I(1)</td><td>I(1)</td></tr>
<tr><td></td><td>$e+p^{T*}$</td><td>I(1)</td><td>I(1)</td></tr>
<tr><td>S1</td><td></td><td>FMOLS</td><td>DOLS</td></tr>
<tr><td></td><td></td><td>1.19
[19.28]</td><td>1.09
[14.51]</td></tr>
<tr><td>S2</td><td></td><td colspan="2">F-Statistics from Wald Test</td></tr>
<tr><td></td><td></td><td>3.06
(0.00)</td><td>1.24
(0.23)</td></tr>
<tr><td colspan="4"></td></tr>
<tr><td colspan="4"></td></tr>
<tr><td rowspan="7">Thailand</td><td>Unit Root Testing</td><td>Variables</td><td>ADF</td><td>DF-GLS</td><td rowspan="7">No</td></tr>
<tr><td></td><td>p^T</td><td>I(1)</td><td>I(1)</td></tr>
<tr><td></td><td>$e+p^{T*}$</td><td>I(1)</td><td>I(1)</td></tr>
<tr><td>S1</td><td></td><td>FMOLS</td><td>DOLS</td></tr>
<tr><td></td><td></td><td>1.32
[13.62]</td><td>1.24
[12.48]</td></tr>
<tr><td>S2</td><td></td><td colspan="2">F-Statistics from Wald Test</td></tr>
<tr><td></td><td></td><td>3.34
(0.00)</td><td>2.43
(0.02)</td></tr>
<tr><td rowspan="7">Hong Kong_1</td><td>Unit Root Testing</td><td>Variables</td><td>ADF</td><td>DF-GLS</td><td rowspan="7">No</td></tr>
<tr><td></td><td>p^T</td><td>Greater than I(1)</td><td>Greater than I(1)</td></tr>
<tr><td></td><td>$e+p^{T*}$</td><td>I(0)</td><td>I(1)</td></tr>
<tr><td>S1</td><td></td><td>FMOLS</td><td>DOLS</td></tr>
<tr><td></td><td></td><td>1.72
[4.75]</td><td>3.07
[3.59]</td></tr>
<tr><td>S2</td><td></td><td colspan="2">F-Statistics from Wald Test</td></tr>
<tr><td></td><td></td><td>2.00
(0.06)</td><td>2.42
(0.02)</td></tr>
<tr><td rowspan="7">Hong Kong_2</td><td>Unit Root Testing</td><td>Variables</td><td>ADF</td><td>DF-GLS</td><td rowspan="7">No</td></tr>
<tr><td></td><td>p^T</td><td>Greater than I(1)</td><td>Greater than I(1)</td></tr>
<tr><td></td><td>$e+p^{T*}$</td><td>I(0)</td><td>I(1)</td></tr>
<tr><td>S1</td><td></td><td>FMOLS</td><td>DOLS</td></tr>
<tr><td></td><td></td><td>1.81
[4.11]</td><td>3.43
[3.23]</td></tr>
<tr><td>S2</td><td></td><td colspan="2">F-Statistics from Wald Test</td></tr>
<tr><td></td><td></td><td>1.84
(0.08)</td><td>2.29
(0.03)</td></tr>
</table>

Country	Verification of Model Conditions				Does PPP Hold for Tradables Prices?
Hong Kong_3	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	Mixed
		p^T	Greater than I(1)	Greater than I(1)	
		$e + p^{T*}$	I(0)	I(1)	
	S1		**FMOLS**	**DOLS**	
			1.53 [4.60]	2.61 [3.17]	
	S2		**F-Statistics from Wald Test**		
			1.59 (0.12)	1.96 (0.06)	
Hong Kong_4	**Unit Root Testing**	**Variables**	**ADF**	**DF-GLS**	No
		p^T	Greater than I(1)	Greater than I(1)	
		$e + p^{T*}$	I(0)	I(1)	
	S1		**FMOLS**	**DOLS**	
			1.96 [4.34]	3.74 [3.55]	
	S2		**F-Statistics from Wald Test**		
			2.13 (0.04)	2.60 (0.02)	

Korea, Pakistan, Philippines and Thailand are four countries for which I reject the existence of PPP with the U.S. For these countries, the traded sector prices hold a significant cointegrating relationship with corresponding U.S. prices, as evidenced by both FMOLS and DOLS model estimates. Further, when the long-run coefficients from the two cointegration regression estimators are tested for equality with one, I reject the null hypothesis and conclude that the coefficient is significantly different from one. As a result, I conclude that the tradables price gap of these countries with the U.S. is a plausible determinant of their long-run real exchange rate appreciation.

The results for Singapore and Sri Lanka produce mixed findings with respect to PPP. The FMOLS and DOLS estimates support the existence of long-run co-movement between home and U.S. traded sector prices. However, only the DOLS estimates fail to reject the null of equi-proportional, long-run co-movement between home and U.S. prices. The FMOLS estimates reject this null, and thus are evidence against PPP. As a result, I will proceed by using the modified model to test the BS hypothesis.

Hong Kong is somewhat complicated because there are four sectoral division schemes. For three of the four, there is strong evidence against PPP. However, for the third scheme of sectoral division (HKG_3), while there is evidence of long-run co-movement between home and U.S. prices, there is mixed evidence in favour of proportional, long-run co-movement. This is sufficient for me to use the modified model for all four sectoral divisions when testing the BS hypothesis for Hong Kong.

Indonesia and Malaysia are the only two countries for which I find strong support for PPP. Their traded sector prices are found to co-move with U.S. prices. Further, inter-country tradables prices show evidence of an equi-proportionate, long-run relationship, as the Wald tests fail to reject the hypothesis that the respective coefficients equal one. Hence, I conclude that there is no statistical basis for including the tradables price gap for the two countries with the U.S. in the real exchange rate equation. I will thus not use the modified model to test the BS hypothesis for these two countries.

Japan deserves special mention because it is the only country in my sample that fails the test of PPP assumptions at its initial stage (S1). FMOLS and DOLS test results suggest that that country's traded sector prices do not converge to U.S. prices. This implies that Japanese prices are capable of retaining their individuality in international markets even in the long-run. These results are in line with Ito et al. (1999), who find sustained departures in Japan's traded sector prices vis-à-vis the U.S. This behaviour of Japan's tradables prices makes the country an ideal candidate to use the modified version of the BS model.

I proceed by estimating the modified model of the BS hypothesis for each country individually, with the exceptions of Indonesia and Malaysia. As these two countries showed evidence of PPP with the U.S., and as the BS hypothesis was previously tested with this assumption for these countries, we do not include them in the subsequent analysis.

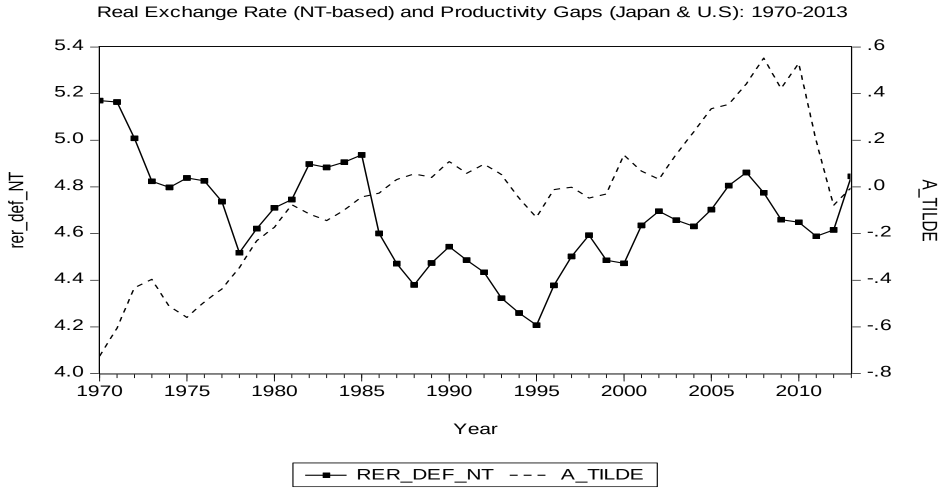
8.3.2 Japan

Starting with Japan, the time plots of rer_def_nt and $\tilde{a}$ (against the U.S.) are given in FIGURE 8.1. Prior to 1995, the rer_def_nt series trended downwards, a pattern compatible with BS propositions. However, after this time, the series trends upward. The $\tilde{a}$ series tends to rise throughout the sample period. The infrequent intersection of the two series suggests that it is unlikely that they have a long-run relationship.

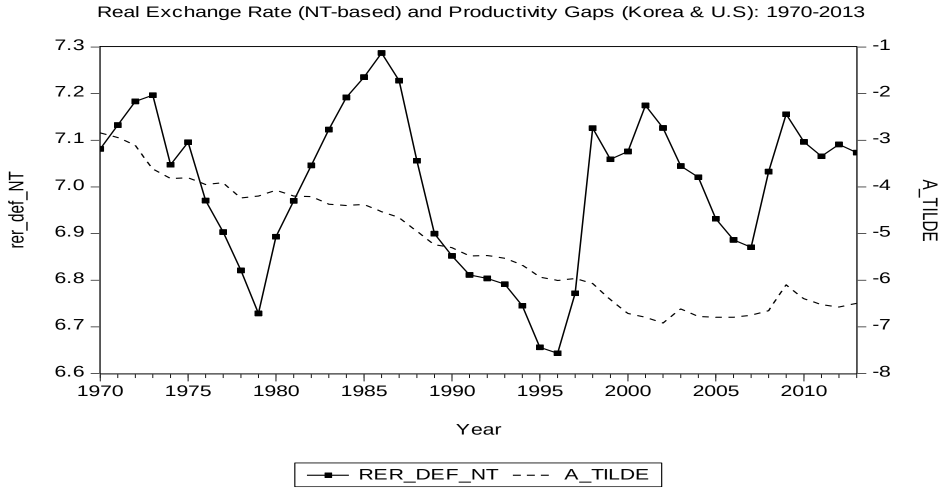
Next, I formally test the rer_def_nt series of Japan and U.S. against rer_def_t and $\tilde{a}$ as model regressors. To recap the test procedures used in earlier chapters, both single equation and multivariate cointegration procedures are used to test long term cointegrating relationships between model variables. If a cointegrating relationship is established for the single equation models, long-run relationships are estimated using FMOLS and DOLS. For the multivariate case, the long-run relationship is estimated by the cointegrating equation as part of the VEC model.

The tests results are reported in TABLE 8.3 below. ADF and DF-GLS tests indicate that all three model time-series are unit root processes.

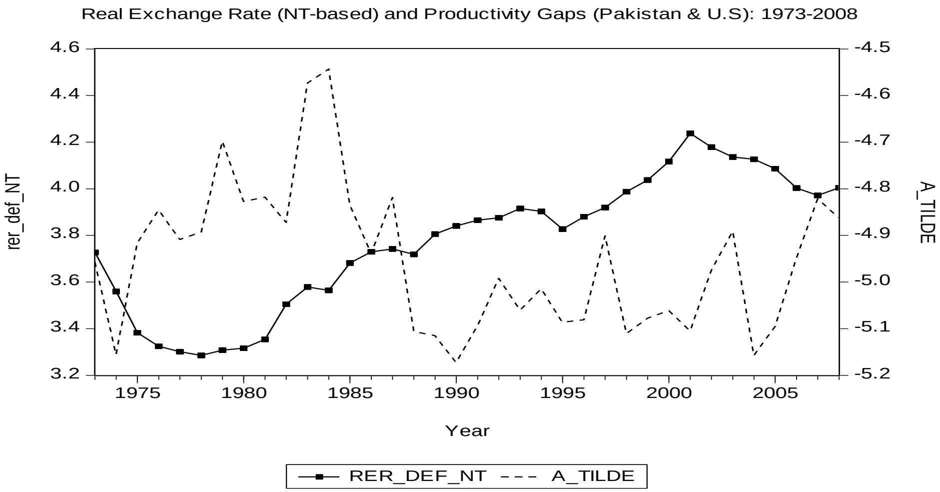
FIGURE 8.1: Plots for Inter Country Relative Prices (Nontradables) and Productivity Differentials against U.S.



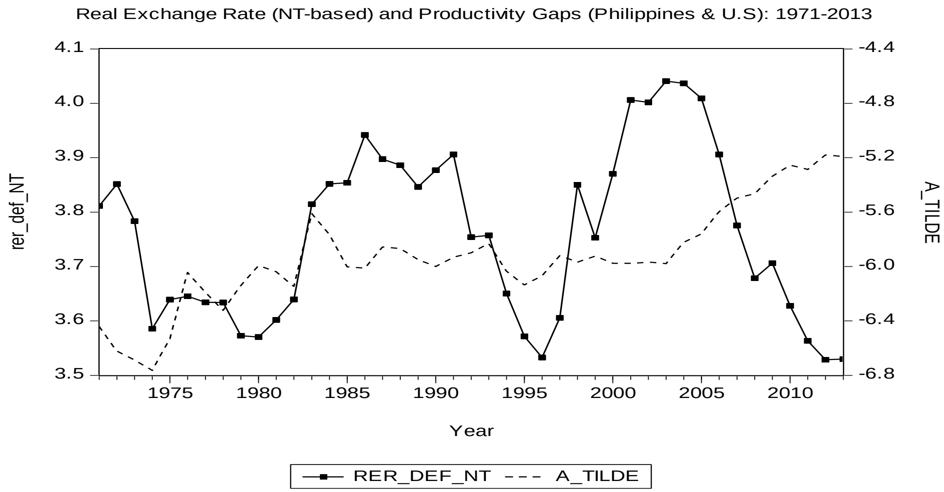
Japan



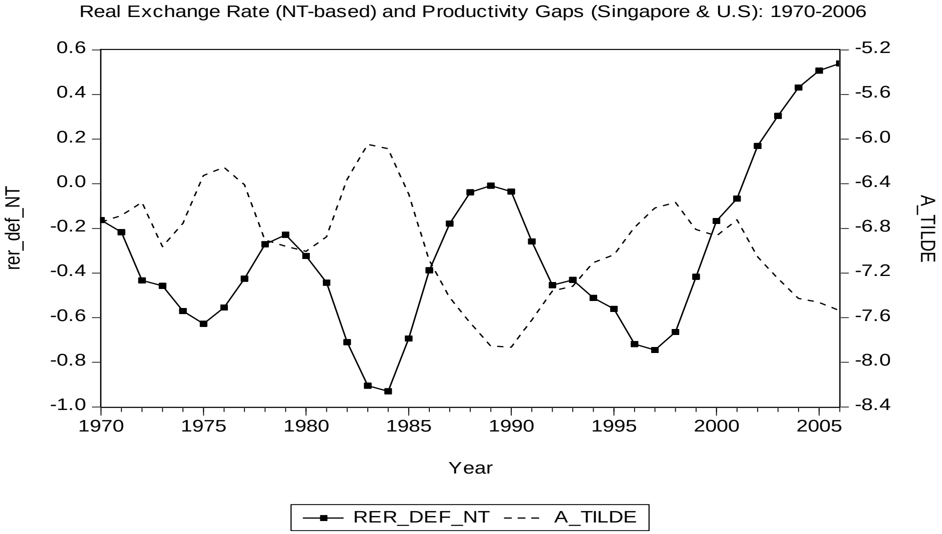
Korea



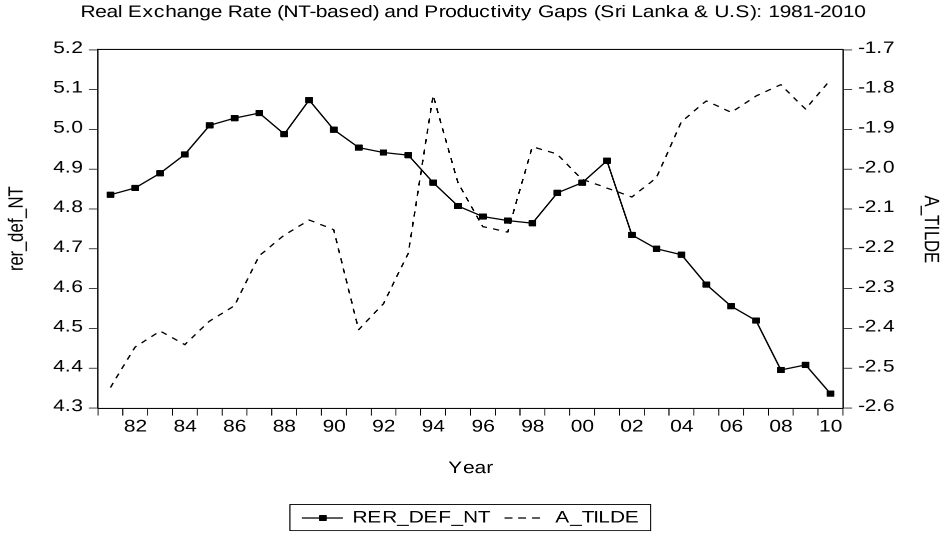
Pakistan



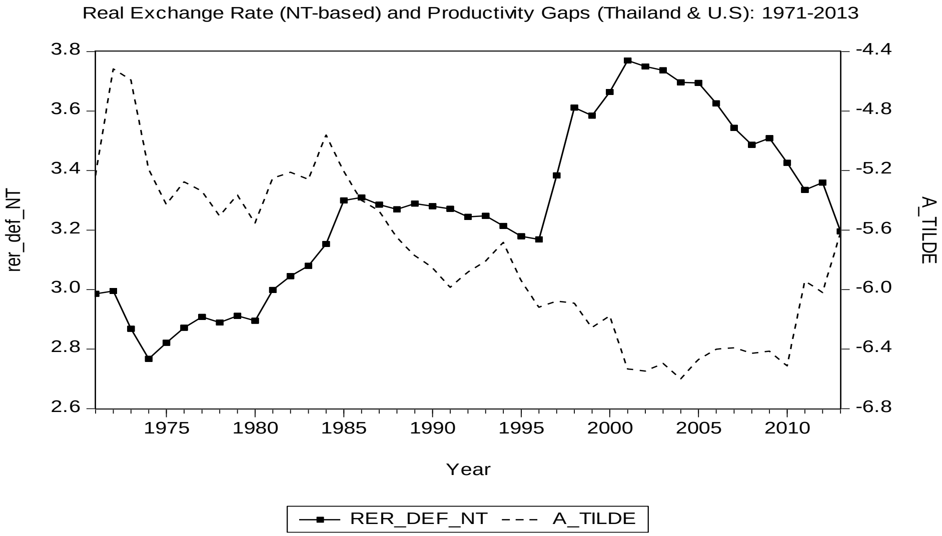
Philippines



Singapore



Sri Lanka



Thailand

TABLE 8.3: Cointegration Tests Results for Japan (1970-2013)[131]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[132]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	1	-0.02 [-0.09]	$\tilde{a}$	-	-	No
				rer_def_t	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

131 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
132 t-values are given in squared-brackets.

Discussing the single equation cointegration models first, neither of the two tests for cointegration find evidence of long-run co-movement between model variables. The residuals from the EG test are concluded to be nonstationary, and the error correction term in the single equation EC model is statistically insignificant. Thus, single equation cointegration tests do not support the existence of a BS effect for Japan.

The multivariate cointegration method also yields no support for the BS hypothesis. The test results are reported in the third panel of TABLE 8.3. Both the Trace statistic and Maximum Eigenvalue statistic for both Cases 3 and 4 find no evidence of a cointegrating vector. In every instance, the tests indicate a rank of zero, suggesting that the series are not cointegrated. As a result, we conclude this as evidence against the BS hypothesis.

8.3.3 Korea

A visual inspection of the time-series plots for rer_def_nt and $\tilde{a}$ for Korea against the U.S. provide weak, informal evidence that a long-run relationship exists between these variables. The two series intersect at several points, with the exchange rate series providing the adjustments, consistent with the existence of a BS effect.

TABLE 8.4: Cointegration Tests Results for Korea (1970-2013)[133]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[134]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	2	-0.27 [-2.26]	$\tilde{a}$	-0.11 [-1.90]	-0.09 [-1.06]	No
				rer_def_t	0.63 [2.57]	0.63 [1.49]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	1	1	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

Vector Error Correction Model

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Lags	White Noise Residuals	$EC_{rer\_def\_nt}$	rer_def_t	BS Coefficient	Does BS Effect Hold?
rer_def_nt	1	Yes	0.00 [1.20]	66.31 [3.43]	-7.51 [-1.77]	No

133 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
134 t-values are given in squared-brackets.

There is mixed evidence regarding the existence of a long-run relationship between rer_def_nt and $\tilde{a}$ in TABLE 8.4. While the EG test leads to the conclusion of no cointegration, the EC model produces an error correction coefficient of -0.27 that is statistically significant, and hence evidence in favour of cointegration. The point estimate indicates that in each period, 27 percent of the total deviations of rer_def_nt from long-run equilibrium are adjusted by the series itself, producing a tendency to converge to a steady-state relationship. However, the estimated long-run relationship is not consistent with the BS hypothesis. Both FMOLS and DOLS produce negative and/or statistically insignificant long-run BS coefficients.

The results obtained through the multivariate model are generally similar. There is partial support in favour of cointegration between model variables. Both the Trace and Maximum Eigenvalue statistics of the Johansen cointegration test in Case 3 indicate the existence of a valid cointegrating vector. As a result, I proceed to estimate the VEC model for this case.

However, there are two problems. First, the the error correction coefficient ($EC_{rer\_def\_nt}$) is statistically insignificant. This indicates that the real exchange rate does not adjust to deviations from long-run equilibrium with productivity differentials. Further, the BS coefficient, estimated in the cointegrating equation, is negative and statistically significant. For both reasons, I conclude that the BS effect does not hold for Korea

If I look into the long-run coefficients of rer_def_t, generated through FMOLS and DOLS estimation in the single equation framework, the results are mixed. The DOLS test results indicate an insignificant role for rer_def_t in explaining trend movement in rer_def_nt. However, FMOLS estimates a statistically significant coefficient for rer_def_t, with a value of 0.63. This suggests that the tradeable series plays a significant role in influencing the rer_def_nt series in the long-run. Similar results hold for the VEC model estimates. The coefficient on the rer_def_t

coefficient is large and significant. However, even with allowing for this role of tradable price differences between the home and U.S. markets, I still find no support for the BS hypothesis for Korea.

8.3.4 Pakistan

A visual inspection of the rer_def_nt and $\tilde{a}$ series of Pakistan, given in FIGURE 8.1, allows for the possibility of a long-run association between the two variables, given that the series show some signs of moving towards each other. The two series are not trending in a common direction. However, the price series shows aggressive trend movements over the sample period, revealing its potential to contribute significant corrections to short-term deviations, so that long-run equilibrium of the series can be restored.

Turning to TABLE 8.5, we see that the single equation (EG) test does not provide evidence of a long-run relationship between rer_def_nt and $\tilde{a}$. In contrast, the ECM test results suggest significant adjustments in rer_def_nt to correct shot-run misalignments, responsible for returning the series to long-run equilibrium. The subsequent BS effect estimates generate mixed support for the hypothesis. The FMOLS estimate rejects the possibility of a valid long-run co-movement between model variables, whereas DOLS results support the hypothesis. Thus, there is empirical ambivalence about the presence of a BS effect for Pakistan.

TABLE 8.5: Cointegration Tests Results for Pakistan (1973-2008)[135]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)*	I(1)
rer_def_t	Yes	I(1)***	I(1)**	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[136]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	3	-0.09 [-1.73]	$\tilde{a}$	-0.14 [-0.32]	2.43 [2.16]	Mixed
				rer_def_t	1.59 [3.46]	3.67 [3.91]	

Multivariate Cointegration Approach

<u>*Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR*</u>

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	1	1	See below

<u>*Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR*</u>

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	1	0	See below

Vector Error Correction Model

	Lags	White Noise Residuals	$EC_{rer\_def\_nt}$	rer_def_t	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_def_nt	2	Yes	-0.00 [-0.33]	10.36 [5.80]	10.76 [5.17]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_def_nt	2	Yes	-0.08 [-1.13]	2.38 [7.18]	1.94 [5.56]	No

[135] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[136] t-values are given in squared-brackets.

The multivariate cointegration model findings are rather encouraging. The Trace statistics of Case 3 and 4 and the Maximum Eigenvalue of Case 3 (only) of the Johansen ML test produce a rank of one, indicating the presence of one valid cointegrating vector between model variables. However, the VEC model does not produce evidence to support the existence of a BS effect for Pakistan. Under both specification 3 and 4 of the test, the model pre-condition is not met, i.e., the test produces a statistically insignificant EC coefficient. This implies that the trend movements of rer_def_nt are not adjusting to restore the series to long-run equilibrium.

Noticeably, the coefficient value of rer_def_t is always large and statistically significant, as shown by both single equation and multivariate cointegration tests. Thus changes in the rer_def_t series appear to be the driver of long-run trend movements of rer_def_nt. Summarizing the results across both single equation and multivariate frameworks, there is 'Mixed' empirical evidence in support of the BS hypothesis for Pakistan and the U.S., but only after controlling for the absence of PPP between traded sector prices of the two countries.

8.3.5 Philippines

The model time-series for the Philippines are plotted in FIGURE 8.1. The country's rer_def_nt and $\tilde{a}$ series do not exactly trend in a common direction. However, their frequent intersection makes their long-run association plausible.

TABLE 8.6: Cointegration Tests Results for Philippines (1971-2013)[137]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[138]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC [139] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	2	-0.20 [-2.04]	$\tilde{a}$	0.37 [2.60]	0.58 [2.46]	Yes
				rer_def_t	1.03 [3.17]	1.27 [2.31]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

[137] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[138] t-values are given in squared-brackets.
[139] The test regression also contain first and second lagged-differences of rer_def_t and $\tilde{a}$.

The empirical estimates for Philippines are reported in TABLE 8.6. Overall, there is sufficient empirical evidence in support of BS effect existing for the country. Looking at the EG single equation cointegration test results, the residuals obtained through regressing rer_def_nt on two of its determinants are not displaying a mean reverting behaviour, necessary to ensure a valid long-run co-movement in model variables. However, the ECM yields a valid error correction coefficient for rer_def_nt, where the series demonstrates an adjustment to short-term fluctuations at a rate of 20 percent per period to restore its long-run equilibrium. When tested for BS coefficient, both FMOLS and DOLS estimators produce positive and statistically significant long-run slope coefficients, confirming the BS effect for the country.

The results obtained from the multivariate model do not indicate cointegration amongst the model variables. The Trace and Maximum Eigenvalue statistics for both cases of the Johansen ML cointegration test (Case 3 and Case 4) do not find a valid cointegrating vector(s) for the estimated models. Both model specifications produce a rank of zero, indicating the inexistence of a long-run association between productivity differential and relative price movements. Thus, based on the rank test estimates, I conclude that the BS effect does not hold valid for the Philippines.

The role of rer_def_t in driving rer_def_nt trend movements is evident from both single equation cointegration regression estimators (FMOLS and DOLS). rer_def_t is associated with significant and sizeable changes to rer_def_nt, as shown by the respective estimated coefficients. This provides evidence on large and persistent deviations of tradables prices of Philippines and the U.S from long-run PPP based equilibrium.

8.3.6 Singapore

Relative to other countries, the time plots for Singapore are not very revealing. The two series (rer_def_nt and $\tilde{a}$ between Singapore and U.S.) does not display a clear time trend. For large part of the sample period, both the series fluctuate around their natural mean values. The latter suggests that the series may be cointegrated, but the lack of a common time trend indicates that there has not been substantial changes to long-run equilibrium over time.

The first evidence against the series being cointegrated is provided by the single equation tests. The EG test is unable to reject the null hypothesis of no cointegration. The ECM also yields a statistically insignificant EC coefficient, and holds a theoretically incorrect sign. Thus, the model variables fail to evidence cointegration both in terms of their residuals as well as the respective error correction process. This lack of co-movement in the long-run behaviour of the series argues against the existence of a BS effect for Singapore.

The multivariate cointegration model also is unable to detect a valid long-run association between the model variables. I obtain a rank of zero according to the Trace and Maximum Eigenvalue statistics, under both specifications (Cases 3 and 4) of the Johansen ML cointegration test. This provides evidence against the series being cointegrated.

Putting together the results from the single equation and multivariate models, I conclude that the BS effect does not exist for Singapore.

TABLE 8.7: Cointegration Tests Results for Singapore (1970-2006)[140]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(0)**	I(0)**	I(0)
rer_def_t	Yes	I(0)**	I(0)**	I(0)
$\tilde{a}$	Yes	I(1)**	I(0)**	Inconclusive

Single Equation Cointegration Approach[141]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC [142] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	2	0.05 [1.15]	$\tilde{a}$	-	-	No
				rer_def_t	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

[140] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[141] t-values are given in squared-brackets.
[142] The test regressions also contain first lagged-differences of rer_def_t and $\tilde{a}$.

8.3.7 Sri Lanka

The visual inspection of the time-series plots for Sri Lanka display a lack of support for the BS effect. This is because (a) the two series are trending in opposite directions for a large part of the sample period and, (b) the rare intersection of two series makes their long-run association implausible.

TABLE 8.8 reports the results of the different statistical tests of the BS hypothesis. The single equation model finds no support for a cointegrating relationship. The tau-statistic, generated by the EG test is unable to reject the null hypothesis of no cointegration. Similarly, the ECM generates an invalid and statistically insignificant error correction process. Taken together, I conclude that the BS effect does not hold for Sri Lanka, on the basis of the two single equation cointegration tests.

There is somewhat more evidence of a cointegrating relationship using the multivariate cointegration framework. The Trace statistic from Case 4 of the Johansen ML test finds support for the existence of a valid cointegrating vector. This is not confirmed by the Maximum Eigenvalue test, nor by either of the tests using the Case 3 specification. Nevertheless, this is enough for me to proceed to estimate short- to long-run dynamics of the model under Case 4 of the VEC model specifications.

The VEC model results do not support the existence of a causal effect between $\tilde{\alpha}$ and rer_def_nt for Sri Lanka and the U.S. in the desired manner. The underlying pre-condition of the model is not met, i.e., short-lived fluctuations in rer_def_nt are not significantly corrected by systematic movements of the series. Also, the long-run BS coefficient bears a negative sign and is statistically significant, implying a depreciating effect (instead of appreciating) of relative sectoral productivity bias of tradables on rer_def_nt. Such a relationship for rer_def_nt and $\tilde{\alpha}$ runs counter to what would be expected for the BS effect. As a result, I reject the BS hypothesis for Sri Lanka.

Similar to previously analysed countries, I find evidence that rer_def_t contributes to significant movements in $\_def\_nt$, driving the series away from its long-run equilibrium.

8.3.8 Thailand

A visual inspection of rer_def_nt and $\tilde{a}$ for Thailand, given in FIGURE 8.1, gives little indication of a long-run association between the two variables, as they largely trend in different directions. That being said, the two series separate and then come together at the end of the sample period, which at least raises the possibility that they could be cointegrated.

TABLE 8.9 reports the results of empirically testing the BS hypothesis for Thailand. Turning first to the single equation EG cointegration test results, there are no signs of long-run cointegration between rer_def_nt and its long-run determinants. The ECM test results are no different from those of the EG test findings. The test produces a statistically insignificant EC coefficient, implying deficient adjustments made by rer_def_nt to return to long-run equilibrium.

Unlike the single equation cointegration models, the Johansen cointegration test finds support for the existence of cointegration for rer_def_nt and $\tilde{a}$, at least for Case 3. The Trace and Maximum Eigenvalue statistics for this specification indicate that there exists a single cointegrating vector. However, upon further estimation, I do not find evidence to support the BS hypothesis. The error correction term in the rer_def_nt equation is insignificant. Further, $\tilde{a}$ is estimated to induce depreciation in rer_def_nt, instead of appreciation as predicted by the BS hypothesis. Thus, in line with the single equation cointegration tests, the multivariate model indicates the inexistence of a BS effect for Thailand.

TABLE 8.8: Cointegration Tests Results for Sri Lanka (1981-2010)[143]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	I(0)***	I(0)

Single Equation Cointegration Approach[144]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	1	-0.07 [-0.26]	$\tilde{a}$	-	-	No
				rer_def_t	-	-	

Multivariate Cointegration Approach

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	0	0	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	1	0	See below

Vector Error Correction Model

	Lags	White Noise Residuals	$EC_{rer\_def\_nt}$	rer_def_t	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_def_nt	0	Yes	-0.05 [-0.85]	1.09 [2.28]	-1.93 [-4.10]	No

[143] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[144] t-values are given in squared-brackets.

TABLE 8.9: Cointegration Tests Results for Thailand (1971-2013)[145]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive

Single Equation Cointegration Approach[146]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC [147] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
rer_def_nt	No	1	-0.09 [-1.12]	$\tilde{a}$	-	-	No
				rer_def_t	-	-	

Multivariate Cointegration Approach

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	0	0	No

Vector Error Correction Model

	Lags	White Noise Residuals	$EC_{rer\_def\_nt}$	rer_def_t	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
rer_def_nt	0	Yes	-0.02 [-1.50]	6.21 [5.29]	-2.22 [-4.31]	No

[145] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[146] t-values are given in squared-brackets.
[147] The test regressions also contain first and second lagged-differences of rer_def_t and $\tilde{a}$.

8.3.9 Summary of Individual Country Studies

In my continued effort to find evidence for the BS hypothesis, I empirically examined the modified version of the BS model in the preceding sections of this chapter. A summary report of individual country estimates is provided in TABLE 8.10. In addition to the inter-country sectoral productivity differential ($\tilde{a}$), I allow for tradables prices of home and U.S. (rer_def_t) to deviate from their PPP equilibrium, and thus to contribute to the trend behaviour of nontradables prices based real exchange rate (rer_def_nt). Nevertheless, this effort proved fruitless. The results are little changed from those of preceding chapters, which investigated the international version of the BS model under standard theoretical settings.

Of all the countries I examined, only Pakistan and the Phillipines showed any evidence of a BS effect, though even here the evidence was mixed, with supporting evidence coming only from the single equation cointegration models. As a result, I conclude that relaxing the assumption of the BS model about PPP for inter-country traded sector prices does not bring substantive changes to my earlier findings.

However, I did find ample evidence that PPP does not hold for most of the countries I am examining. For eight out of ten sample countries, home and foreign traded sector prices were found to display sustained departures from long-run PPP. rer_def_t was generally found to have a significant, long-run relationship with rer_def_nt. As a result, this series establishes itself as an important driver of real exchange rates. These findings are in agreement with empirical results of earlier studies on Asia (Ito et al., 1999; Thomas and King, 2008).

TABLE 8.10: Does Modified Balassa-Samuelson Hypothesis Hold? Summary of Results by Country

Country	**Individual Results**			**Country Summary**
	Single Equation Cointegration Method	*Multivariate Cointegration Method*		
		Case 3	*Case 4*	
Japan (1970-2013)	No	No	No	No
Korea (1970-2013)	No	No	No	No
Pakistan (1973-2008)	Mixed	No	No	Mixed
Philippines (1971-2013)	Yes	No	No	Mixed
Singapore (1970-2006)	No	No	No	No
Sri Lanka (1981-2010)	No	No	No	No
Thailand (1971-2013)	No	No	No	No

8.3.10 Hong Kong

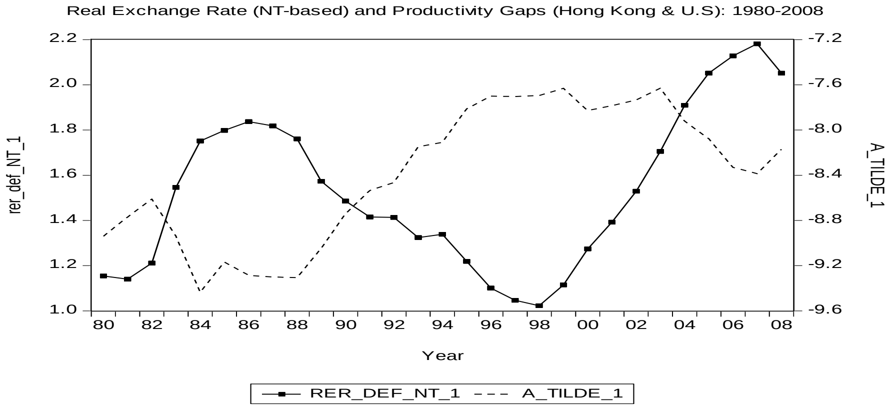
I conduct a separate country study for Hong Kong because of issues associated with matching appropriate sectoral classifications for the country (see Chapter 3, Section 3.3.3 for a detailed discussion on sectoral classifications for Hong Kong). As previously, I will analyse four separate pairs of time series, one for each type of sectoral division.

FIGURE 8.2 plots time series of rer_def_nt and $\tilde{a}$ for all four sectoral divisions. Overall, the visual evidence seems lacking to support the hypothesis of a long-run association between these variables. For almost the entire data period, the two series move in a dissimilar directions (except for HKG_2). Further, their infrequent intersection points are not indicative of an adjustment process whereby the price series adjust to productivity differences to return to long-run equilibrium.

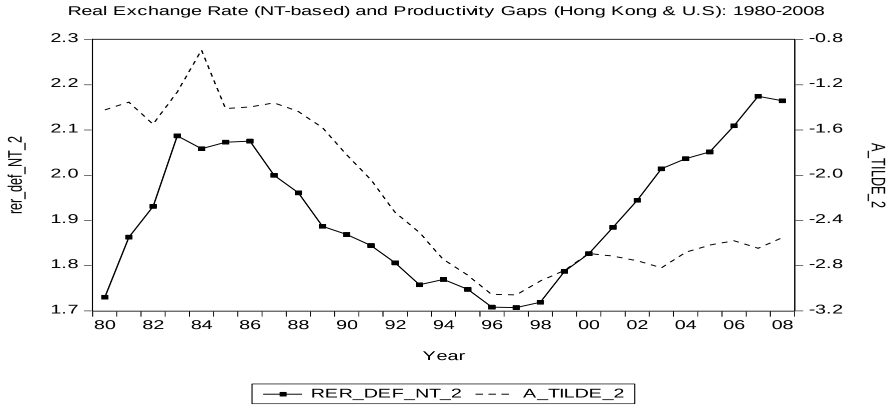
TABLE 8.11 reports the results of cointegration analysis for the first sectoral division, HKG_1. According to the EG test, there is no significant long-run association between rer_def_nt and $\tilde{a}$. The associated tau statistic, when compared against MacKinnon (1996) critical values, does not reject the null hypothesis of no cointegration at the 10 percent significance level. In contrast, the ECM does find evidence, as given by a statistically significant error correction term. However, the subsequent long-run slope coefficients of the BS model, as estimated by FMOLS and DOLS, show the wrong (negative) sign. Thus, I conclude, on the basis of the single equation cointegration model, that the BS effect does not hold for HKG_1.

A somewhat different picture is provided by the multivariate cointegration framework. There is strong evidence from both cointegration tests (Trace and Maximum Eigenvalue) and for both cases (Case 3 and 4) of a single cointegrating vector. As a result, I proceed with estimation of the VEC model for both cases/specifications.

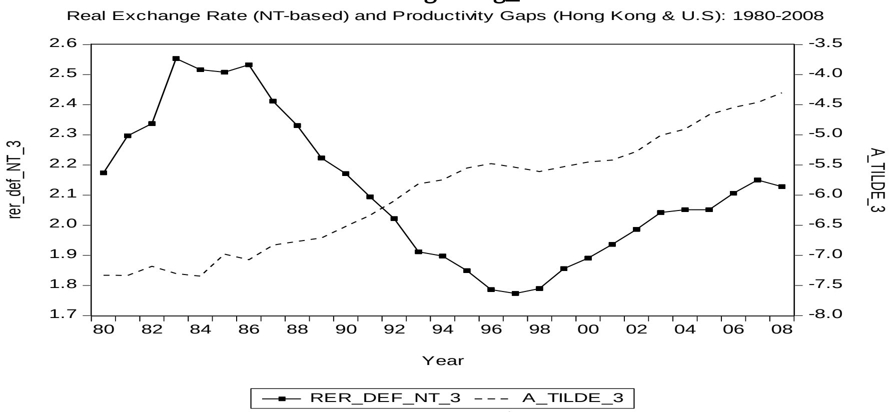
FIGURE 8.2: Plots for Relative Prices (Nontradables) and Productivity Differentials (Hong Kong: 1980-2008)



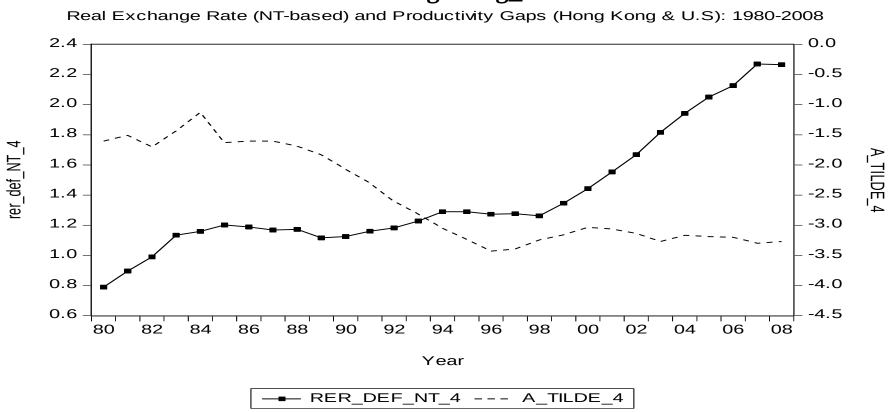
Hong Kong_1[148]



Hong Kong_2[149]



Hong Kong_3[150]



Hong Kong_4[151]

[148] Construction is the only nontradable sector here. Rest of all the sectors are treated as tradables.
[149] Mining, utilities, construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.
[150] Mining, utilities and construction are nontradable sectors whereas rest of all the sectors are treated as tradables.
[151] Construction and wholesale & retail trade are nontradable sectors whereas rest of all the sectors are treated as tradables.

TABLE 8.11: Cointegration Tests Results for Hong Kong: Case 1 (1980-2008)[152]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_1$	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_t\_1$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_1$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[153]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC [154] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
$rer\_def\_nt\_1$	No	2	-0.13 [-1.94]	$\tilde{a}\_1$	-0.79 [-1.78]	-1.93 [-2.38]	No
				$rer\_def\_t\_1$	-1.64 [-1.48]	-4.36 [-2.16]	

Multivariate Cointegration Approach

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
$rer\_def\_nt\_1$	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
$rer\_def\_nt\_1$	1	1	See below

Vector Error Correction Model

Dependent Variable	Lags	White Noise Residuals	$EC_{rer\_def\_nt\_1}$	$rer\_def\_t\_1$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
$rer\_def\_nt\_1$	1	Yes	0.05 [2.07]	-13.96 [-4.57]	-5.93 [-4.76]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$rer\_def\_nt\_1$	1	Yes	-0.18 [-2.79]	6.89 [4.94]	1.62 [3.61]	Yes

[152] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[153] t-values are given in squared-brackets.
[154] The test regression also contains first lagged-difference of $rer\_def\_t\_1$ and $\tilde{a}\_1$.

The two specifications of the VEC model (Cases 3 and 4) produce contrasting results. Case 3 of the model does not favour the valid existence of a BS effect as the model pre-condition is not successfully met. The error correction term in the $rer\_def\_nt\_1$ equation is positive and significant, which is not consistent with a series that responds to deviations by returning to long-run equilibrium. Further, the BS coefficient is negative and statistically significant, which is opposite of what the BS hypothesis predicts. In contrast, specification 4 of the model is supportive of the BS hypothesis. Both the EC coefficient and long-run BS coefficient are statistically significant and have the right signs, negative and positive, respectively.

Both specifications of VEC model estimate statistically significant long-run slope coefficients for $rer\_def\_t\_1$. Similar to previously discussed countries, this indicates PPP does not hold between the country's and U.S.'s tradable prices. This runs counter to the classical formulation of the BS theory, which assumes no role for tradables prices in displacing real exchange rates from their long-run equilibrium.

Overall, I conclude that there is mixed evidence in support of the BS hypothesis for Hong Kong when I use the first sectoral division for the respective series.

TABLE 8.12: Cointegration Tests Results for Hong Kong: Case 2 (1980-2008)[155]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_2$	Yes	I(1)**	I(1)**	I(1)
$rer\_def\_t\_2$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_2$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[156]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC[157] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
$rer\_def\_nt\_2$	No	2	-0.08 [-0.99]	$\tilde{a}\_2$	-	-	No
				$rer\_def\_t\_2$	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_2$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_2$	1	0	See below

Vector Error Correction Model

Dependent Variable	Lags	White Noise Residuals	$EC_{rer\_def\_nt\_2}$	$rer\_def\_t\_2$	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
rer_def_nt_2	1	Yes	0.66 [2.00]	1.16 [16.09]	-0.12 [-3.99]	No

155 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
156 t-values are given in squared-brackets.
157 The test regression also contains first lagged-difference of $rer\_def\_t\_2$ and $\tilde{a}\_2$.

I briefly summarize the analysis of the remaining three sectoral divisions. Sectoral divisions 2 and 4 show no support for the BS hypothesis in either single equation or multivariate cointegration frameworks. There is mixed support using the third sectoral division (HKG_3). While the EG test rejects the existence of a long-run association between *rer_def_nt_3* and $\tilde{a}$_3, the EC model produces a negative and statistically significant error correction term, suggesting relatively quick adjustment to deviations from long-run equilibrium. Subsequent estimation of the long-run relationship using FMOLS and DOLS finds positive and statistically significant BS coefficients, consistent with the BS hypothesis.

With respect to rer_def_t, I once again find that this series is frequently a significant determinant of rer_def_nt. The VECM results for sectoral division 2 (Case 4), the single equation FMOLS and DOLS results for sectoral division 3, the VECM results for sectoral division 3 (Case 4), and the VECM results for sectoral division 4 (both Cases 3 and 4), all indicate that rer_def_t is positively and significantly related to rer_def_nt. Thus, I reject the PPP assumption in the classical formulation of the BS model, which assumes no significant role for inter-country traded sector prices in determining the long-run behaviour of the real exchange rate.

TABLE 8.13: Cointegration Tests Results for Hong Kong: Case 3 (1980-2008)[158]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_3$	Yes	I(1)**	I(1)***	I(1)
$rer\_def\_t\_3$	Yes	I(1)**	I(0)**	Inconclusive
$\tilde{a}\_3$	Yes	I(0)**	I(0)***	I(0)

Single Equation Cointegration Approach[159]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC[160] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
$rer\_def\_nt\_3$	No	2	-0.51 [-1.93]	$\tilde{a}\_3$	0.04 [2.19]	0.05 [1.88]	Yes
				$rer\_def\_t\_3$	1.49 [14.40]	1.39 [11.36]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_3$	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_3$	1	1	See below

Vector Error Correction Model

Dependent Variable	Lags	White Noise Residuals	$EC_{rer\_def\_nt\_3}$	$rer\_def\_t\_3$	BS Coefficient	Does BS Effect Hold?
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$rer\_def\_nt\_3$	1	Yes	0.14 [0.28]	1.29 [30.41]	-0.13 [-3.68]	No

[158] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[159] t-values are given in squared-brackets.
[160] The test regression also contains first lagged-difference of $rer\_def\_t\_3$ and $\tilde{a}\_3$.

TABLE 8.14: Cointegration Tests Results for Hong Kong: Case 4 (1980-2008)[161]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_4$	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_t\_4$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_4$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[162]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC[163] Coeff	Variables	LR Coefficient FMOLS	LR Coefficient DOLS	Does BS Effect Hold?
$rer\_def\_nt\_4$	No	2	-0.01 [0.25]	$\tilde{a}\_4$	-	-	No
				$rer\_def\_t\_4$	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_4$	1	1	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_4$	1	2	See below

Vector Error Correction Model

Dependent Variable	Lags	White Noise Residuals	$EC_{rer\_def\_nt\_4}$	$rer\_def\_t\_4$	BS Coefficient	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
$rer\_def\_nt\_4$	1	Yes	0.10 [3.32]	6.56 [6.39]	-2.65 [-6.70]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
$rer\_def\_nt\_4$	1	Yes	0.43 [3.48]	2.55 [12.51]	-0.74 [-7.92]	No

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161 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
162 t-values are given in squared-brackets.
163 The test regression also contains first lagged-difference of $rer\_def\_t\_4$ and $\tilde{a}\_4$.

8.4 Panel Data Estimation Results

I start my panel data analysis by formally testing the model variables using a mix of panel unit root and stationarity tests. As before, I test for unit roots using the Fisher-ADF, Fisher-PP and Hadri tests.

I will first test the model variables involved in the assumption of PPP; i.e., p^T and $e + p^T$. For home prices of tradables (p^T), the two unit root tests indicate that the series are level stationary. In contrast, the Hadri stationarity test supports a conclusion that the variable is integrated of some higher order (greater than order one, I(1)). As the two unit root tests "outweigh" the single stationarity test, I conclude p^T to be level-stationary. For U.S prices of tradables, converted to the home country's unit of currency $(e + p^T)$, all three tests produce common results. Accordingly, I conclude that this variable is integrated of order one.

With respect to testing PPP for inter-country traded sector prices, I follow the protocol established in Section 8.5.1 for the individual country analysis. For pooled data, the existence (inexistence) of PPP will be tested against conditions P1 and P2, respectively; which are simply the panel analogues to the single country conditions S1 and S2.

There is mixed evidence in support of equi-proportionate long-run co-movement between home and U.S tradables prices. PFMOLS and PDOLS both estimate significant long-run co-movement between home and U.S. traded sector prices (P1). However, I obtain inconsistent results when testing P2. The test based on PFMOLS suggests a disproportionate long-run relationship between home and foreign prices, whereas the test based on PDOLS fails to reject a one-to-one movement between the two prices. The fact that one of the tests rejects PPP is sufficient for me to proceed by estimating the modified (panel) BS model.

Before performing cointegration tests on the modified BS model, I first seek evidence on the order of integration of the three model variables (rer_def_nt, rer_def_t, and $\tilde{a}$). The results are reported at the top of TABLE 8.15. With respect to rer_def_nt, the two unit root tests produce contrasting results. The variable is unit root in levels according to the Fisher-ADF test, whereas the Fisher-PP test and the Hadri stationarity tests indicate the series to be integrated of order one. Thus, I conclude the series to be a unit root process in levels. For rer_def_t, all three tests consistently show that the series is stationary in differences. Finally, with respect to $\tilde{a}$, the two unit root tests conclude that the series is level stationary, while the Hadri stationarity test indicates an order of integration greater than 1.

The third and fourth panels of TABLE 8.15 display the test results for the Pedroni and Johansen Fisher panel cointegration tests. I first discuss the Pedroni tests. When performing the Pedroni cointegration tests, I opted for automatic lag selection. All seven statistics associated with the suite of Pedroni cointegration tests consistently failed to reject the null hypothesis of no cointegration between model variables at the 10 percent significance level. Based on the Pedroni tests, I conclude that the series are not cointegrated and thus reject the BS hypothesis for the data set of nine developing Asian economies.

Unlike the Pedroni tests, the Fisher-Johansen panel cointegration tests support the existence of a long-run association between model variables. The Trace and Maximum Eigenvalue statistics of the two specifications (Case 3 and Case 4) indicate a rank of one, suggesting one cointegrating vector for the estimated model.

TABLE 8.15: Summary of Results for Panel Unit Root and Cointegration Tests for Modified Balassa-Samuelson Hypothesis[164,165]

Panel Unit Root Test Results (Order of Integration as Determined by)				
Variables	Fisher-ADF	Fisher-PP	Hadri	Conclusion
p^T	I(0)***	I(0)***	Greater than I(1)	I(0)
$e + p^T$	I(1)***	I(1)***	I(1)***	I(1)
rer_def_nt	I (0)*	I (1)***	I (1)***	I (1)
rer_def_t	I (1)***	I (1)***	I (1)***	I(1)
$\tilde{a}$	I (0)**	I (0)**	Greater than I (1)	I(0)

Testing for Tradables PPP[166]				
P1. Testing for Cointegration		**P2. Testing for Equi-Proportionate Relationship** ***(F-Statistics from Wald Coefficient Test)***		**Does PPP Hold for Tradables Prices?**
PFMOLS	PDOLS	PFMOLS	PDOLS	Mixed
1.08 [37.73]	1.01 [34.49]	7.25 (0.01)	0.06 (0.80)	

[164] ***, ** and * are representing significance of sample statistics at 1%, 5% and 10% levels respectively.
[165] Hong Kong is omitted from panel estimations.
[166] t-ratios and p-values are given in squared-brackets and parenthesis, respectively.

Estimating the Modified Version of the Balassa-Samuelson Hypothesis							
Pedroni Panel Cointegration Test Results[167]							
Common AR Coefficients (Within Dimension)				*Individual AR Coefficients (Between Dimension)*			
Panel v Statistics	Panel ρ Statistics	Panel PP Statistics	Panel ADF Statistics	Group ρ Statistics	Group PP Statistics	Group ADF Statistics	Does BS Effect Hold?
-0.13	1.96	2.10	1.33	2.00	1.30	0.65	No

Johansen-Fisher Panel Cointegration Test Results[168,169]

Case 3: Intercept (no trend) in cointegrating equation and VAR

Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigenvalue)	Does BS Effect Hold?
1	1	See below

Case 4: Intercept and trend in cointegrating equation-no trend in VAR

Fisher Stat (From Trace Stat)	Fisher Stat (From Max-Eigenvalue)	Does BS Effect Hold?
1	1	See below

[167] Pedroni panel cointegration is a test for null of no cointegration in both homogenous and heterogeneous panels. The test statistics are standardized and asymptotically normally distributed. See Pedroni (1995, 1999) for further details.
[168] The test is maximum likelihood based rank test.
[169] Lag selection is done through SIC under panel VAR.

Long-Run Cointegrating Vectors for the Modified Balassa-Samuelson Hypothesis
Results for Panel FMOLS and DOLS[170] Estimators

Estimator	Long-Run Coefficient[171]		Does BS Effect Hold?
	rer_def_t	$\tilde{a} = BS\ Coefficient$	
PFMOLS	0.30 [2.70]	-0.13 [-2.87]	No
PDOLS	0.27 [1.88]	-0.12 [-2.24]	No

[170] Lead = Lag = 1,
[171] t-values are given in squared-brackets.

Having established the possibility of cointegration from the Fisher-Johansen panel cointegration test results, I proceed to estimate the long-run relationship between rer_def_nt, on the one hand, and rer_def_t and $\tilde{a}$, using panel FMOLS (PFMOLS) and panel DOLS (PDOLS) single equation cointegration regression estimators. Both panel estimators yield a negative and statistically significant long-run BS coefficient. This is inconsistent with the prediction of the BS hypothesis. As a result, I interpret this as strong evidence against the BS hypothesis.

Finally, in the context of long-run PPP between inter-country tradables prices, the long-run elasticities produced by the PFMOLS and PDOLS estimators suggest a significant contribution of rer_def_t to the deviations of rer_def_nt from its long-run equilibrium. Thus, lack of PPP between home and U.S. tradables generates trend departures in the home country's real exchange rate. These results are consistent with previous results regarding violation of the assumption of tradables PPP built into the classical formulation of the BS hypothesis.

8.4.1 Summary of Panel Data Results

This chapter has found strong evidence against the assumption of PPP that is incorporated in the classical version of the BS model. However, allowing for the divergence of tradables prices from PPP does not bring us any closer to finding support for the BS hypothesis. The associated estimates from the pooled data analysis strongly reject the modified version of the BS model. These findings are in line with my individual country results as well as earlier analyses of the model (in preceding chapters) using the standard (international) version. Thus, if we are to find support for the BS hypothesis, we must look elsewhere. The next chapter relaxes further assumptions built into the BS model in the hope that by doing so, we may uncover evidence of a BS effect for the subject Asian economies.

TABLE 8.16: Does the Modified Balassa-Samuelson Hypothesis Hold? Summary of Results for Panel Cointegration Tests

Test of Cointegration			
Pedroni Residual Based Panel Cointegration Test	Johansen-Fisher Panel Cointegration Test		Conclusion
	Case 3	Case 4	
No	No	No	No

APPENDIX-D

EViews Programming Code for Korea

```
wfopen    "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-8\Korea.wf1"
'*********************************************************************
'Group Plots for Real Exchange Rate (NT) & Productivity Gap
'*********************************************************************
group gA rer_def_NT A_tilde
freeze(group_plot) gA.line(x)
group_plot.setelem(1) lcolor(black) symbol(7) lpat(1)
group_plot.setelem(2) lcolor(black) symbol(4) lpat(1)
group_plot.setelem(3) lcolor(black) symbol(1) lpat(1)
group_plot.setelem(3) lcolor(black)
group_plot.options linepat
group_plot.addtext(t) Real Exchange Rate (NT-based) and Productivity Gaps (Korea & U.S): 1970-2013
group_plot.addtext(b) Year
group_plot.addtext(l) PNT
group_plot.addtext(l) NER
group_plot.addtext(l) rer_def_NT
group_plot.addtext(r) A_TILDE
'***********************************************************
'***********************************************************

create y 1970 2013
'importing data from Excel for Korea
import  "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-8\Chapter 8.xlsx" range="Korea"

'***************************************************************************************************
'VERIFYING THE ASSUMPTION OF PPP FOR Korea AND U.S. TRADABLES PRICES
'***************************************************************************************************
'************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'************************************************************

'*************************
'Graph for Korea's  pT
'*************************

genr pT = pT
freeze(figure_pT) pT.line
figure_pT.addtext(t) pT (Korea):  1970-2013
figure_pT.addtext(b) Year
figure_pT.addtext(l) pT
figure_pT.legend(off)

'We see from the FIGURE that pT has a time trend to it.  So we would include both an intercept and a time trend in our unit root regression equtions.

'***************************************
'ADF Unit Root Test for Korea's pT
'***************************************

'We now run our first ADF test

freeze(table_8_1_1_pT_adf) pT.uroot(adf,trend,lag=1)

'Note that I select lags, p = 1 for obtaining white residuals.  The unit root test produces a t-value of -4.90 which is smaller than our 5% criterion -3.52.  Thus, at this point, we can reject the null of a unit root.
```

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the adf test and finally will check for white noise.

```
genr resid = 0
freeze(mode=overwrite,pT_adf) pT.uroot(adf,const,trend,lag=1)
freeze(pT_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the pT series is level stationary.

'***
'DF-GLS Unit Root Test for Korea's pT
'***

```
freeze(table_8_1_3_pT_dfgls) pT.uroot(dfgls,trend,lag=1)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.36 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

```
genr pTdiff = d(pT)
freeze(table_8_1_4_pTdiff1_dfgls) pTdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -1.96 which is smaller than our 5% criterion -1.95. Thus, we may now reject the null of non-stationarity in first differenced series of pT.

''Putting it all together, I conclude that the pT series is I(1), a finding incompatible with my ADF test results.

'*************************************
'Graph for Korea's pT_us_PPP
'*************************************

```
genr pT_us_PPP = pT_us_PPP
freeze(figure_pT_us_PPP) pT_us_PPP.line
figure_pT_us_PPP.addtext(t) pT_us_PPP (Korea):  1970-2013
figure_pT_us_PPP.addtext(b) Year
figure_pT_us_PPP.addtext(l) pT_us_PPP
figure_pT_us_PPP.legend(off)
```

'We see from the FIGURE that pT_us_PPP has a time trend to it. So we would include both an intercepT_us_PPP and a time trend in our unit root regression equtions.

'**
'ADF Unit Root Test for Korea's pT_us_PPP
'**

'We now run our ADF test

```
freeze(table_8_1_1_pT_us_PPP_adf) pT_us_PPP.uroot(adf,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -2.01 which is greater than our 5% criterion -3.52. Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the adf test and finally will check for white noise.

```
genr resid = 0
freeze(mode=overwrite,pT_us_PPP_adf) pT_us_PPP.uroot(adf,const,trend,info=sic)
freeze(pT_us_PPP_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the pT_us_PPP series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr pT_us_PPPdiff = d(pT_us_PPP)
freeze(figure_pT_us_PPPdiff) pT_us_PPPdiff.line
figure_pT_us_PPPdiff.addtext(t) DpT_us_PPP (Korea):  1970-2013
figure_pT_us_PPPdiff.addtext(b) Year
figure_pT_us_PPPdiff.addtext(l) DpT_us_PPP
figure_pT_us_PPPdiff.legend(off)
```

'The graph is not particularly illuminating. Depending on how you look at it, it could have a time trend to it. However, from the graphical inspection of the series, I conclude that it does not have a time trend, we run the ADF test with a constant only.

'So we begin the whole process over again:

```
genr pT_us_PPPdiff = d(pT_us_PPP)
freeze(table_8_1_2_pT_us_PPPdiff1_adf) pT_us_PPPdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -4.73 which is smaller our 5% criterion -2.93. Thus, we may now reject the null of non-stationarity in first differenced series of pT_us_PPP. There is no reason to go further. The last thing we do is check for white noise.

```
genr resid = 0
freeze(mode=overwrite,pT_us_PPPdiff1_adf) pT_us_PPPdiff.uroot(adf,const,trend,info=sic)
freeze(pT_us_PPPdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the pT_us_PPP series is I(1).

'**
'DF-GLS Unit Root Test for Korea's pT_us_PPP
'**

```
freeze(table_8_1_3_pT_us_PPP_dfgls) pT_us_PPP.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.09 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

```
genr pT_us_PPPdiff = d(pT_us_PPP)
freeze(table_8_1_4_pT_us_PPPdiff1_dfgls) pT_us_PPPdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -4.37 which is smaller than our 5% criterion -1.95. Thus, we may now reject the null of non-stationarity in first differenced series of pT_us_PPP.

''Putting it all together, I conclude that the pT_us_PPP series is I(1), a finding compatible with my ADF test results.

''**
''S1: Establishing Cointegration for Verifying PPP for Inter-Country Tradables Prices
'**

'Now, by employing FMOLS and DOLS cointegration regression estimators, we shall calculate our LR cointegrating vectors.

```
equation table_8_1_LRPPP_fmols.cointreg(method=fmols) pT pT_us_PPP

equation table_8_1_LRPPP_dols.cointreg(method=dols, trend=constant, lag=1,lead=1 ) pT pT_us_PPP
```

'As proven by FMOLS and DOLS test results, valid cointegration holds between pT and pT_us_PPP. Thus, I can proceed with estimation of S2 condition.

```
'******************************************************************************************************
'S2: Testing for Equi-proportionate Relationship between Inter-Country Tradables Prices
'******************************************************************************************************

freeze(table_8_1_LRPPP_fmols_Ftest) table_8_1_LRPPP_fmols.wald c(1)=1

freeze(table_8_1_LRPPP_dols_Ftest) table_8_1_LRPPP_dols.wald c(1)=1
'As proven by both Wald coefficient test results for FMOLS and DOLS estimates, PPP does not validly hols for traded sector prices of Korea and U.S. Thus, I can proceed with estimating the modified version of BS hypothesis.

'****************************************************************************************************
'ESTIMATING BALASSA-SAMUELSON EFFECT FOR rer_def_NT, rer_def_T  & a_tilde
'****************************************************************************************************
'************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'************************************************************
'**************************************************************
'Graph for Korea's  rer_def_NT
'**************************************************************

genr rer_def_NT = rer_def_NT
freeze(figure_rer_def_NT) rer_def_NT.line
figure_rer_def_NT.addtext(t) rer_def_NT (Korea):  1970-2013
figure_rer_def_NT.addtext(b) Year
figure_rer_def_NT.addtext(l) rer_def_NT
figure_rer_def_NT.legend(off)

'We see from the FIGURE that rer_def_NT has a time trend to it.  So we would include both an intercept and a time trend in our unit root regression equtions.

'*****************************************************
'ADF Unit Root Test for Korea's RER_DEF_NT
'*****************************************************

freeze(table_8_3_1_rer_def_nt_adf) rer_def_nt.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -2.93 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,rer_def_nt_adf) rer_def_nt.uroot(adf,const,trend,info=sic)
freeze(rer_def_nt_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the rer_def_nt series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr rer_def_ntdiff = d(rer_def_nt)
freeze(figure_rer_def_ntdiff) rer_def_ntdiff.line
figure_rer_def_ntdiff.addtext(t) drer_def_nt (Korea):  1970-2013
figure_rer_def_ntdiff.addtext(b) Year
figure_rer_def_ntdiff.addtext(l) drer_def_nt
figure_rer_def_ntdiff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_8_3_2_rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -4.73 which is now smaller than our 5% criterion -2.93. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def_nt. There is no reason to go further. The last thing we do is to check ADF regression result for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_def_ntdiff1_adf) rer_def_ntdiff.uroot(adf,const,info=sic)
freeze(rer_def_ntdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_nt series is I(1).

'**
'DF-GLS Unit Root Test for Korea's RER_DEF_NT
'**

```
freeze(table_8_3_3_rer_def_nt_dfgls) rer_def_nt.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -2.97 which is greater than our 5% criterion -3.19. Thus, we may not reject the null of non-stationarity for rer_def_nt.

'Now let's see if the series is difference stationary or not

```
genr rer_def_ntdiff = d(rer_def_nt)
freeze(table_8_3_4_rer_def_ntdiff1_dfgls) rer_def_ntdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -4.63 which is now smaller than our 5% criterion -1.95. Thus, we may reject the null of non-stationarity in first differenced series of a_tilde.

''Putting it all together, I conclude that the rer_def_nt series is I(1), a finding compatible with my ADF test results.

'***********************************
'Graph for Korea's rer_def_T
'***********************************

```
genr rer_def_T = rer_def_T
freeze(figure_rer_def_T) rer_def_T.line
figure_rer_def_T.addtext(t) rer_def_T (Korea): 1970-2013
figure_rer_def_T.addtext(b) Year
figure_rer_def_T.addtext(l) rer_def_T
figure_rer_def_T.legend(off)
```

'We see from the FIGURE that rer_def_T has a time trend to it. So we would include both an intercept and a time trend in our unit root regression equtions.

'**
'ADF Unit Root Test for Korea's rer_def_T
'We now run our first ADF test
'**

```
freeze(table_8_3_1_rer_def_T_adf) rer_def_T.uroot(adf,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 1. The unit root test produces a t-value of -3.12 which is greater than our 5% criterion -3.52. Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise. To do that, I first set all the residuals = 0, then run the adf test and finally will check for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_def_T_adf) rer_def_T.uroot(adf,const,trend,info=sic)
freeze(rer_def_T_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_T series is not level stationary.
'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr rer_def_Tdiff = d(rer_def_T)
```

```
freeze(figure_rer_def_Tdiff) rer_def_Tdiff.line
figure_rer_def_Tdiff.addtext(t) Drer_def_T (Korea):  1970-2013
figure_rer_def_Tdiff.addtext(b) Year
figure_rer_def_Tdiff.addtext(l) Drer_def_T
figure_rer_def_Tdiff.legend(off)
```

'The graph is not particularly illuminating. Depending on how you look at it, it could have a time trend to it. However, from the graphical inspection of the series, I conclude that it does not have a time trend, we run the ADF test with a constant only.

'So we begin the whole process over again:

```
genr rer_def_Tdiff = d(rer_def_T)
freeze(table_8_3_2_rer_def_Tdiff1_adf) rer_def_Tdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =1. The unit root test produces a t-value of -4.91 which is smaller our 5% criterion -2.93. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def_T. There is no reason to go further. The last thing we do is check for white noise.

```
genr resid = 0
freeze(mode=overwrite,rer_def_Tdiff1_adf) rer_def_Tdiff.uroot(adf,const,trend,info=sic)
freeze(rer_def_Tdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the rer_def_T series is I(1).

'***
'DF-GLS Unit Root Test for Korea's rer_def_T
'***

```
freeze(table_8_3_3_rer_def_T_dfgls) rer_def_T.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -1.94 which is greater than our 5% criterion -3.19. Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

```
genr rer_def_Tdiff = d(rer_def_T)
freeze(table_8_3_4_rer_def_Tdiff1_dfgls) rer_def_Tdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -5.44 which is smaller than our 5% criterion -1.95. Thus, we may now reject the null of non-stationarity in first differenced series of rer_def_T.

''Putting it all together, I conclude that the rer_def_T series is I(1), a finding compatible with my ADF test results.

'*******************************
'Graph for Korea's a_tilde
'*******************************

```
genr a_tilde = a_tilde
freeze(figure_a_tilde) a_tilde.line
figure_a_tilde.addtext(t) a_tilde (U.S):  1970-2013
figure_a_tilde.addtext(b) Year
figure_a_tilde.addtext(l) a_tilde
figure_a_tilde.legend(off)
```

'We see from the FIGURE that a_tilde has a time trend to it. So we would include both an intercept and a time trend in our unit root regression equations.

```
'****************************************************
'ADF Unit Root Test for Korea's Productivity
'****************************************************

freeze(table_8_3_1_a_tilde_adf) a_tilde.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.5 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.
genr resid = 0
freeze(mode=overwrite,a_tilde_adf) a_tilde.uroot(adf,const,trend,info=sic)
freeze(a_tilde_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the a_tilde series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr a_tildediff = d(a_tilde)
freeze(figure_a_tildediff) a_tildediff.line
figure_a_tildediff.addtext(t) da_tilde (Korea):  1970-2013
figure_a_tildediff.addtext(b) Year
figure_a_tildediff.addtext(l) da_tilde
figure_a_tildediff.legend(off)

'From the graph, the series clearly does not have a time trend to it. So, I would test the series for unit with an intercept only.

'So we begin the whole process over again:

genr a_tildediff = d(a_tilde)
freeze(table_8_3_2_a_tildediff1_adf) a_tildediff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0.  The unit root test produces a t-value of -5.81 which is now smaller than our 5% criterion -2.93.  Thus, we may now reject the null of non-stationarity in first differenced series of a_tilde.  There is no reason to go further.  The last thing we do is to check ADF regression result for white noise.

genr resid = 0
freeze(mode=overwrite,a_tildediff1_adf) a_tildediff.uroot(adf,const,info=sic)
freeze(a_tildediff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the a_tilde series is I(1).

'*******************************************************
'DF-GLS Unit Root Test for Korea's Productivity
'*******************************************************

freeze(table_8_3_3_a_tilde_dfgls) a_tilde.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.15 which is greater than our 5% criterion -3.19.  Thus, at this point, we may not reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr a_tildediff = d(a_tilde)
freeze(table_8_3_4_a_tildediff1_dfgls) a_tildediff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 0.  The unit root test produces a t-value of -5.88 which is now smaller than our 5% criterion -1.94. Thus, we may reject the null of non-stationarity in first differenced series of a_tilde.

''Putting it all together, I conclude that the a_tilde series is I(1), a finding compatible with my ADF test results.
```

```
'**********************************************
'Single Equation Cointegration Methods
'**********************************************

'*****************************************************************
''Graph the suspected cointegrated series together
'*****************************************************************
'The first step is to print out a graph of the series.  This is very important!

group g1 rer_def_NT rer_def_T a_tilde
freeze(figure1) g1.line(x)
figure1.setelem(1) lcolor(black) symbol(1) lpat(1)
figure1.setelem(2) lcolor(black) symbol(4) lpat(1)
figure1.setelem(3) lcolor(black) symbol(7) lpat(1)
figure1.setelem(3) lcolor(black)
figure1.options linepat
figure1.addtext(t) rer_def_NT, lnpT & a_tilde (Korea & U.S): 1970-2013
figure1.addtext(b) Year
figure1.addtext(l) rer_def_NT
figure1.addtext(l) lnpT
figure1.addtext(r) a_tilde

''******************************************************
''S1.A.Engle-Granger Approach to Cointegration
'******************************************************

genr resid = 0
equation eg.ls rer_def_NT c rer_def_T a_tilde
genr EC1 = resid

'First we test if the residuals of above regression are level stationary or not. If yes, next we'll proceed towards estimation of error correction model.

'***************************
'Graph for Korea's  EC
'***************************
genr EC1 = EC1
freeze(figure_EC1) EC1.line
figure_EC1.addtext(t) EC1 (Korea):  1970-2013
figure_EC1.addtext(b) Year
figure_EC1.addtext(l) EC1
figure_EC1.legend(off)

'*******************************************************************************
'EG Test for Cointegration
'********************************************************************************

freeze(table_8_3_EGC) g1.coint(method=eg)

'The null hypothesis will be accepted as suggested by sample statistics.

''******************************************
''S1.B.Error Correction Model (ECM)
'******************************************
''******************************************************
'Selecting the number of lags in the VAR  *
'******************************************************
'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var var1.ls 1 6   g1
freeze(var1_lagtest1) var1.laglen(6)
freeze(var1_lagtest2) var1.testlags
'The laglength test above indicates that the should have VAR has 1 lag. Bu the residuals are not white noise at 1 lag. So, I have to raise the number of lags to 2.

var var2.ls 1 2  g1
```

```
freeze(var2_artest1) var2.correl
freeze(var2_artest2) var2.qstats(12)
freeze(var2_artest3) var2.arlm(12)

'The residuals are white noise. So I am satisfied with the selection of 2 lags.

'We now try different lags of d(rer_def_T) and d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_def_NT c rer_def_T a_tilde
genr ec1 = resid

var table_8_3_eg2a.ls 0 0 d(rer_def_NT)   @  c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2))

var table_8_3_eg2b.ls 0 0 d(rer_def_NT)  @  c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2)) d(rer_def_T(-1)) d(a_tilde(-
1))

var table_8_3_eg2c.ls 0 0 d(rer_def_NT)   @   c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2)) d(rer_def_T(-1))
d(rer_def_T(-2)) d(a_tilde(-1)) d(a_tilde(-2))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.

var table_8_3_eg2a.ls 0 0 d(rer_def_NT)   @   c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2))
freeze(table_8_3_eg2a_artest1) table_8_3_eg2a.correl
freeze(table_8_3_eg2a_artest2) table_8_3_eg2a.qstats(12)
freeze(table_8_3_eg2a_artest3) table_8_3_eg2a.arlm(12)

'The residuals are not absolutely white noise. Let's try Model B and C.

var table_8_3_eg2b.ls 0 0 d(rer_def_NT)   @   c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2)) d(rer_def_T(-1))
d(a_tilde(-1))
freeze(table_8_3_eg2b_artest1) table_8_3_eg2b.correl
freeze(table_8_3_eg2b_artest2) table_8_3_eg2b.qstats(12)
freeze(table_8_3_eg2b_artest3) table_8_3_eg2b.arlm(12)

var table_8_3_eg2c.ls 0 0 d(rer_def_NT)   @   c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2)) d(rer_def_T(-1))
d(a_tilde(-1)) d(a_tilde(-1)) d(a_tilde(-2))
freeze(table_8_3_eg2c_artest1) table_8_3_eg2c.correl
freeze(table_8_3_eg2c_artest2) table_8_3_eg2c.qstats(12)
freeze(table_8_3_eg2c_artest3) table_8_3_eg2c.arlm(12)

'Neither of the three models yield white residuals. So I prefer to continue my next set of estimatons with Model A.

''************************
''Estimating EC Model
'*************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for
establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_8_3_ecm.ls(n) d(rer_def_NT) c ec1(-1) d(rer_def_NT(-1)) d(rer_def_NT(-2))

'Note that the SR effect is significant as the EC coefficient is of value -0.27 is statistically significant at better than 1%
significance level.
```

''***
''S2.A & S2.B: Obtaining LR Coefficients
'**

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient, i.e., BS coefficient for Korea against U.S.

equation table_8_3_LReqn1_fmols.cointreg(method=fmols) rer_def_NT rer_def_T a_tilde

equation table_8_3_LReqn2_dols.cointreg(method=dols, trend=constant, lag=2,lead=2) rer_def_NT rer_def_T a_tilde

'The BS coefficient obtained through FMOLS and DOLS estimators are -0.11 and -0.09, i.e., the long-run BS coefficients are bearing incorrect signs. Thus, there is invalid evidence in support of BS effect existing for Korea.

''***
'Multivariate Cointegration Approach
'***

''***
''Check if the VAR(2) model is dynamically stable
'***
freeze(table_8_3_var2_varstable) var2.arroots(graph)

'The model is dynamically stable.

''**
''M1.A & M1.B: Identifying the number of cointegrating vectors
'***
'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_8_3_var2_coint) var2.coint(s,2)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. All the results indicate 1 cointegrating vectors.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in Eviews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

''***
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'***

' For estimating the LR relationship, corresponding VEC command is:

var table_8_3_vec_c.ec(c,1) 1 1 rer_def_NT rer_def_T a_tilde

'CONCLUSION: I conclude that rer_def_NT and a_tilde are not cointegrated in the Korea's data.

CHAPTER 9: THE BALASSA-SAMUELSON HYPOTHESIS AND THE DEMAND-SIDE DETERMINANTS OF REAL EXCHANGE RATE

9.1 Motivation

In the preceding chapters, I attempted to find an explanation for trend appreciation in real exchange rates for Asia by establishing its connection with productivity (supply-side) shocks. However, it is well recognized that, in the long-run, real exchange rates can be driven by demand-side factors. These include wealth, consumer preferences, government spending patterns, external balance of the country, demographic and institutional changes, and many more. In addition to sectoral productivity imbalances, the demand-side factors may also significantly determine the long-run deviations in real exchange rates as they evolve over time. Unfortunately, there is no consensus on the most cohesive and comprehensive set of demand-side factors that determine trend departures in real exchange rates from long-run PPP equilibrium.

In Chapter Eight, I investigated the validity of one of the core assumptions of the Balassa-Samuelson (BS) hypothesis; i.e., that PPP holds for cross-country tradables prices. I was unable to find much support for this assumption. The failure of PPP further opens the door for demand-side factors to transmit their effects to real exchange rates (MacDonald, 2007).

The primary objective of this chapter is to investigate if the inclusion of demand-side determinants of real exchange rates within the BS model changes the conclusion from my prior findings, which overwhelmingly rejected the existence of a BS effect for developing Asian states. The next section briefly discusses some of the demand-side factors that have been found responsible (in empirical studies) for causing long-run real exchange rate misalignments. Keeping in view the data availability and study sample period, I will focus on a few of these factors. The selected determinants will then be incorporated in the modified version of the BS model (from Chapter Eight) as I continue to search for evidence of a BS effect for Asia.

9.2 Demand-Side Shocks and Domestic Sectoral Prices

Shifts in demand, influencing domestic sectoral prices through both private and public spending, have been analysed by a number of studies in earlier and recent literature (Dornbusch, 1988; Bergstrand, 1991; Chinn, 2000; Halpern and Wyplosz, 2001; Boreo et al., 2015). Countries in the catching up process of development can experience a substantial rise in income levels. Demand-side factors in such economies may significantly affect relative sectoral prices, in the event of non-homothetic consumer preferences. Normally, consumer preferences are thought to be skewed towards nontradables, so that services behave as superior goods. These biased preferences towards nontradables can have a deleterious effect on overall domestic inflation, causing the real exchange rate of the country to experience trend deviations (appreciation) from its long-run equilibrium.

A sizeable number of studies focus on the role of demand-side factors as drivers of the relative prices of nontradables (De Gregorio, Giovannini and Wolf, 1994; De Gregorio, Giovannini and Krueger, 1994; MacDonald and Ricci, 2001). I briefly discuss some of these factors below.

9.2.1 Net Foreign Assets

In the long-run, economies are required to maintain their inter-temporal budget constraints. Lenders demand the repayment of loans and borrowers are required to pay back these loans. This activity of lenders and borrowers is reflected in current account deficits and surpluses as the respective parties adjust their Net Foreign Assets (NFA) holdings. During the course of these adjustments, an increase in a country's NFA position will serve as a net increase in a country's wealth. Such a rise in wealth levels will transmit the same effects to the macro-economy as a rise in income levels. This increase in wealth will eventually translate into a rise in demand for both tradable and nontradable goods. Nontradables being only produced domestically, this will prompt a rise in the relative prices of nontradables. This rise in nontradables prices will not only tend to appreciate the real exchange rate, but will also stimulate nontradables production at the expense of tradables production (given the long-run supply constraints). As a result, current account deficits will arise. The long-run association between NFA and real exchange rates has been examined empirically by a wide range of studies (see Faruqee, 1995; Gagnon, 1996; Alberola et al., 1997; Lane and Milesi-Ferretti, 2000, 2002).

9.2.2 Government Consumption Expenditures

The size of government consumption expenditures, relative to a country's domestic output, is also one of the fundamentals historically considered to be an important determinant of real exchange rates (see Froot and Rogoff, 1991; De Gregorio, Giovannini and Holger, 1994; Ostry, 1994). Rising levels of government expenditures tend to shift domestic demand towards home nontradables, causing the real exchange rate to appreciate. This over-concentration of public sector consumption on nontradables leads to an appreciation in relative prices of nontradables at home (Rogoff, 1992; Chinn and Johnston, 1996; Chinn, 1999).

Modelling a two-period, small open economy, Frenkel and Razin (1996) empirically investigate two distinct channels through which government spending can impact real exchange rates: the resource-withdrawal and consumption-tilting channels. Regarding the first channel, the effect of government expenditure is analogous to that of a negative supply shock. The effect on private consumption and real exchange rates will depend on the proportion of government consumption spending falling on nontradables versus that falling on tradables (see Edwards, 1989; Galstyan and Lane, 2009 for empirical verification of the first channel). Regarding the second channel, the authors point out that the effect of government expenditure on private consumption levels and the real exchange rate will depend upon the characteristics of the utility function. They highlight the potential importance of complementarity versus substitutability between private consumption and government consumption in utility, which determines how the marginal rate of inter-temporal substitution in utility is influenced by government expenditure levels (see Balvers and Bergstrand, 1997, 2002; Ravn et al., 2007).

9.2.3 Terms of Trade

Terms of Trade (TOT) serve as another important factor displacing the real exchange rate from its long-run equilibrium through its impact on the demand for nontradables. An improvement in TOT will raise national income levels, making exports dearer and imports cheaper. A rise in income levels generates a proportionate rise in the demand for both tradables and nontradables. Such a rise in demand for nontradables bears effects similar to a rise in NFA holdings or any other positive demand-side shock. Tradables prices being determined internationally (for small open economies only), improvement in TOT bears positive effects on the relative price of nontradables. The process of inflating nontradables' prices will be intensified by asymmetric consumer preferences which will shift to domestically produced nontradables (medical care, housing, education, entertainment and leisure, etc.), owing to rising income levels. The two-fold effect of rising TOT will translate into the trend appreciation of real exchange rates from its long-run equilibrium. The wealth effects of terms of trade causing trend departures in real exchange rate from its long-run equilibrium level is empirically verified by many studies (see De Gregorio and Wolf, 1994; Mendoza, 1995; Amano and Norden, 1995; Thomas and King, 2008).

9.2.4 Output Per Capita

Real output per capita is another vital demand-side factor, generating wealth effects and causing shifts in the relative price of nontradables. In the long-run, the link between nontradables price appreciations and rising levels of output per capita can be two-fold; (a) the effect may operate through higher demand for the (constrained) supply of nontradables, and (b) biased preferences (non-homothetic tastes) in favor of nontradables, with income elasticity of demand for nontradables greater than one. In either situation, the long-run coefficient on output per capita is expected to be negatively related to the real exchange rate; i.e., an income rise (fall) causes downward (upward) shifts to real exchange rates. The upward long-run pressures on nontradables prices, exerted by output per capita are empirically confirmed by Bergstrand (1991), Eckstein and Friedman (2011), and Nassif et al. (2011).

9.2.5 Real Oil Prices

Real oil prices is an important non-monetary factor, affecting long-run real exchange rates through its impact on the current account balance and real economic activity at home and in the foreign country. Oil price fluctuations may trigger large shifts in the wealth of nations. Unusual upward or downward swings in oil prices may result in large current account imbalances in both oil exporting as well as oil importing countries, thus bringing trend deviations to real exchange rates from its long-run equilibrium.

The transmission mechanisms through which oil prices have an impact on real economic activity include both supply and demand-side channels. The supply-side effects are related to the fact that crude oil is a basic input to production, and consequently an increase in oil prices lead to a rise in production costs which ultimately puts upward pressure on the overall price level of the country. Oil price changes also entail demand-side effects on consumption and investment. Consumption is affected indirectly through its positive relation with disposable income, triggering inflation at the domestic level, thus causing real exchange rate appreciation. The magnitude of this effect is in turn stronger, if the shock is perceived to be long-lasting.

Furthermore, oil prices have an adverse impact on investment by increasing firms' costs. An extensive number of studies have noted the potential importance of oil price movements in long-run real exchange rate determination (see McGuirk, 1983; Krugman, 1983a,b; Golub, 1983; Rogoff, 1991; Clarida and Gali, 1994; Amano and Norden, 1998a,b; Camarero and Tamarit, 2002; Chen and Chen, 2007).

9.3 Model Specification and Data Sources

Keeping in view data availability and the relative importance of numerous demand-side factors in the context of emerging Asian countries, I propose an *Augmented Version of the Modified BS hypothesis*. For establishing the augmented version of the model from the modified BS hypothesis proposed in Chapter Eight, I select real government consumption spending, real output per capita and real oil prices to account for demand-side influences on real exchange rate movements. A number of BS studies have produced very promising results on the subject, using the said demand-side factors as the determinants of long-run real exchange rates (see Chinn, 2000; Thomas and King, 2008; Badia and Ubiergo, 2014)

The inclusion of demand-side factors in the estimable equation of the modified BS hypothesis makes the model look a little different. Recalling the modified version of the BS hypothesis from Chapter Eight (Equation 8.4):

$$(8.4)' \qquad rer\_def\_nt = \alpha + \vartheta(rer\_def\_t) - \beta\tilde{a} + \mu, \text{ where}$$

$rer\_def\_nt =$ Bilateral (comprising of nontradables prices only) real exchange rate between home and the U.S. Nontradables prices are measured through sectoral VA deflators.

$rer\_def\_t =$ Bilateral (comprising of tradables prices only) real exchange rate between home and the U.S. Tradables prices are measured through sectoral VA deflators. The series accounts for the deviations of tradables prices from their long-run PPP (if there are any).

$\tilde{a}_t =$ Inter-country relative sectoral productivity gap between home and the U.S., capturing the BS effect. Incorporating oil prices, government consumption spending and output per capita as demand-side factors, equation $(8.4)'$can be re-stated as:

(9.1)[172] $rer\_def\_nt = \alpha + \vartheta(rer\_def\_t) - \beta\tilde{a}_t + \varphi_i Z_t + \mu.$

Equation (9.1) is the augmented version of the modified BS hypothesis. Z is a 3×1 vector of demand-side determinants of real exchange rates, comprised of real oil prices (rop), real government consumption expenditures (gov_exp) and real GDP per capita $(gdp\_percapita)$ as a measure of output per capita, $Z = [rop, gov\_exp,\ gdp\_percapita]$. rer_def_nt is posited to be a linear function of Z, where φ_i represents the long-run slope coefficients of the respective variables.

[172] Since rer_def_t $, rop, gov\_exp,\ gdp\_percapita$ are not the main variables of interest, their estimated coefficients are always reported in black, irrespective of their statistical significance (insignificance).

TABLE 9.1: Definitions and Data Sources for Demand-Side Factors for Augmenting Balassa-Samuelson Model

Variables	Definition	Source
Real Oil Prices (rop)	Crude petroleum, average of UK Brent (light), Dubai (medium) and Texas (heavy), equally weighted ($/barrel). The series is constructed by deflating nominal oil price data with the U.S. GDP deflator (constant 2005).	United Nations Conference on Trade and Development (UNCTAD)
Government Expenditures (gov_exp)	General government final consumption expenditures (constant 2005 U.S.D). This includes all government current expenditures for purchases of goods and services (including compensation of employees). It also includes most expenditures on national defence and security, but excludes government military expenditures that are part of government capital formation. Data are in constant 2005 U.S. dollars.	World Bank national accounts data and OECD National Accounts data files
Output/GDP per capita ($gdp\_percapita$)	GDP per capita (constant 2005 U.S.D) = gross domestic product divided by midyear population. GDP is the sum of gross value added by all resident producers in the economy plus any product taxes and minus any subsidies not included in the value of the products. It is calculated without making deductions for depreciation of fabricated assets or for depletion and degradation of natural resources. Data are in constant 2005 U.S. dollars.	World Bank national accounts data and OECD National Accounts data files

9.4 Country Studies

Similar to previous chapters, this section will discuss country-by-country analysis of the augmented version of the modified BS hypothesis. The model is re-defined to incorporate demand-side determinants of real exchange rates besides accounting for inter-country productivity differences and the plausible absence of PPP between inter-country tradables prices. The augmented version of the model is tested empirically using equation (9.1), established in Section 9.3 of this chapter. The results are reported by adopting the same econometric procedures as in the preceding chapters. The step-by-step description of econometric methods employed are given through EViews program of Korea, attached in the Appendix to this chapter.

9.4.1 Indonesia

Following on the results from Chapter Eight, I begin by assuming that PPP holds for Indonesia, as I could not reject PPP for Indonesia (and Malaysia). I proceed by implementing tests for establishing long-run co-movement between the rer_def_nt of Indonesia (against the U.S.) and its proposed long-run determinants. To restate, the same empirical tests used in earlier chapters, i.e., both single equation and multivariate cointegration procedures, are used to estimate the respective relationships.

TABLE 9.2 reports the country results for Indonesia assuming PPP. ADF and DF-GLS unit root tests generally indicate that the model variables follow a unit root process. However, the evidence is mixed for $\tilde{a}$. This leaves me uncertain about the true order of integration for this variable.

TABLE 9.2: Cointegration Tests Results for Indonesia[173] (1976-2013)[174]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	Greater than I(1)	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapit$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[175]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.34 [-1.66]

Estimator	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	-0.30 [-2.16]	-1.75 [-3.15]	0.07 [0.20]	0.24 [0.94]	No
DOLS	-0.46 [-1.17]	-0.33 [-0.19]	0.73 [1.07]	-0.18 [-0.35]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	0	0	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	1	1	See below

Vector Error Correction Model[176]

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	$EC_{rer\_def\_nt}$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
rer_def_nt	0.04 [0.62]	-1.11 [-6.77]	-3.14 [-5.54]	-6.38 [-6.85]	2.06 [9.65]	No

[173] The country does not include rer_def_t as a model regressor, since PPP holds for its traded sector prices against U.S. (see Chapter 8).

[174] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.

[175] t-values are given in squared-brackets.

[176] The test regression takes no lagged-differences of model variables.

Turning to the single equation cointegration tests results, the Engle-Granger (EG) residuals indicate a unit root process. Further, the error correction term indicates mean-reversion adjustments to deviations from long-run equilibrium (obtained through regressing rer_def_nt on its determinants). These results provide evidence that the model variables are cointegrated. As a result, I proceed further with FMOLS and DOLS estimation of the model. However, both sets of estimates produce a statistically insignificant, long-run BS coefficient. Accordingly, I conclude that the single equation cointegration tests do not support the existence of a BS effect for Indonesia.

The test results for the multivariate model are reported in the third panel of TABLE 9.2. The Trace and Maximum Eigenvalue statistics detect the presence of a valid cointegrating vector for Case 4 (but not Case 3). This allows me to conduct VEC estimations for this case (see the last panel of the table).

The estimated results are not favourable to the BS hypothesis. Analogous to Chapter Eight, I base my decision about the validity of the BS effect only on the sign and statistical significance of the error correction (EC) coefficient for rer_def_nt. For establishing a valid BS effect, it is necessary that this coefficient be negative and statistically significant. This condition is not met for the VEC model, as the EC coefficient is insignificant. As in the case of the single equation test results, the evidence does not support the existence of a BS effect.

With respect to the demand-side factors, there is evidence from the multivariate estimates that rop, gov_exp and $gdp\_percapita$ induce appreciation in rer_def_nt. However, this finding receives only limited support in the single equation estimates. The multivariate results are in line with earlier literature, revealing negative effects of government spending shocks to real exchange rates. Thus, a rise in government spending serves as a positive wealth effect, subsequently causing positive shifts to private consumption and real wages. As a consequence, domestic nontraded sector relative prices receive an upward push, making the country's real exchange rate appreciate from its long-run equilibrium (Fat´as and Mihov, 2001; Blanchard and Perotti, 2002). Similarly, the positive wealth effects of rising income/output per capita, inducing real exchange rate trend appreciation have received recognition in the empirical literature (Kravis et al., 1982; Obstfeld and Rogoff, 1996).

TABLE 9.2. reports country estimates for Indonesia assuming PPP. Despite failing to reject PPP for Indonesia, I will also re-estimate the model relaxing this assumption. I do this for two reasons. First, as a general practice, failure to reject a hypothesis should not be taken as acceptance of that hypothesis. As a result, an argument can be made for estimating the model under the alternative hypothesis. .

There is another reason for relaxing the assumption of PPP even though my previous analysis could not reject this. Indonesia and Malaysia are the only two countries from my set of 10 Asian states for which I was unable to reject PPP for traded sector prices. In order to include these two countries in my subsequent panel analysis, I need to use the same variable specification across countries. That is, the real exchange rate needs to be modelled using rer_def_t, $\tilde{a}$ and demand-side real factors as regressors.

On the basis of these two reasons, I re-estimate the BS model for Indonesia, this time including the variable rer_def_t as an additional determinant of long-run real exchange rates. The results are reported in TABLE 9.3 .

TABLE 9.3: Cointegration Tests Results for Indonesia (1976-2013)[177]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	Greater than I(1)	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[178]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.34 [-1.66]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?	
FMOLS	1.68 [8.88]	0.16 [2.22]	-0.23 [-0.83]	0.77 [5.35]	0.16 [1.50]	No	Mixed
DOLS	2.28 [10.10]	0.45 [3.40]	-0.69 [-1.56]	0.65 [3.67]	0.36 [2.46]	Yes	

Multivariate Cointegration Approach

	Trace	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	2	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	3	1	See below

Vector Error Correction Model[179]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.14 [1.15]	2.38 [14.40]	0.57 [8.65]	-1.30 [-5.03]	0.37 [2.69]	0.39 [5.40]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
0.17 [1.23]	2.05 [12.02]	0.36 [4.38]	-1.31 [-5.58]	-0.31 [-0.81]	0.48 [6.68]	No

[177] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[178] t-values are given in squared-brackets.
[179] The test regressions take no lagged-differences of model variables.

It turns out that the inclusion of rer_def_t as a regressor in the BS model somewhat alters the conclusions from the preceding analysis. While the EG test does not indicate that the variables are cointegrated, the EC coefficient suggests otherwise. I find that the respective coefficient is negative, of plausible size and significant at the 10 percent level. Upon subsequent estimation of the FMOLS and DOLS equations, I obtain mixed results. The FMOLS test produces a statistically insignificant BS coefficient. However, the DOLS estimator produces a positive long-run slope coefficient which is statistically significant at the 1 percent level.

In contrast, the multivariate cointegration analysis accords with the previous analysis that imposed PPP on estimation of the BS model. Both Trace and Max Eigenvalue tests produce evidence of cointegration among model variables for both cases (Case 3 and Case 4). However, in subsequent estimation of the VEC model, the respective error correction coefficients are insignificant, implying insignificant adjustments on the part of rer_def_nt to adjust to deviations from long-run equilibrium. As a result, I conclude that the multivariate cointegration test results do not support the hypothesis of a BS effect for Indonesia. Taken together, the single equation and multivariate cointegration results produce mixed results in favour of the BS hypothesis.

9.4.2 Japan

Starting with the single equation EG cointegration test results, there is no evidence of long-run cointegration between rer_def_nt and the other, long-run determinants in the augmented, modified version of the BS model. This holds true for both the EG residual test and the EC coefficient. The two tests are unable to reject the null hypothesis of no cointegration. Thus, single equation cointegration methods yield no support for a BS effect in Japan.

TABLE 9.4: Cointegration Tests Results for Japan (1970-2013)[180]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[181]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	0.09 [0.19]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	-	-	-	-	-	No
DOLS	-	-	-	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	2	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	2	See below

Vector Error Correction Model[182]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.32 [-1.92]	1.40 [19.00]	-0.03 [-1.34]	0.40 [3.60]	0.72 [6.16]	-0.03 [-0.38]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.16 [-1.73]	1.65 [12.53]	0.07 [1.79]	1.23 [3.19]	1.86 [6.24]	-0.05 [-0.40]	No

[180] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[181] t-values are given in squared-brackets.
[182] The test regressions take no lagged-differences of model variables.

The multivariate cointegration results are at variance with the single equation results. Both the Trace and Maximum Eigenvalue statistics for Case 3 and Case 4 indicate the existence of two cointegrating vectors; and, thus, that the respective model variables are cointegrated.

Proceeding to the VECM results, I find evidence of a negative and statistically significant error correction term in the rer_def_nt equation. This fulfils a necessary condition for the existence of a BS effect. However, the resulting BS coefficient is insignificant and of the wrong sign. Therefore, I conclude that appreciation in Japan's rer_def_nt is not significantly driven by $\tilde{a}$. In summary, there is no evidence in favour of a BS effect for Japan from either the single equation or multivariate cointegration test results.

In contrast with the empirical findings from Indonesia, the demand-side determinants gov_exp and $gdp\_percapita$ demonstrate a positive long-run association with rer_def_nt, suggesting real exchange rate depreciation. These findings are consistent with the *deep habit model* proposed by Ravn et al. (2006). That model suggests that rising government spending may cause a sizeable fall in mark-ups at home (relative to foreign markets). Hence, the domestic economy becomes less expensive relative to the foreign market, or, equivalently, the real exchange rate depreciates. The depreciating effects of government spending on real exchange rates, according to the deep habit mechanism, is empirically confirmed elsewhere by Ravn et al. (2007).

9.4.3 Korea

The country estimates for Korea are reported in TABLE 9.5. The EG residual test does not find evidence of a cointegrating relation. These results differ from the ECM which does identify a stationary, long relationship between rer_def_nt and its supply and demand-side determinants.

TABLE 9.5: Cointegration Tests Results for Korea (1970-2013)[183]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[184]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.59 [-3.60]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	1.37 [8.07]	0.20 [3.60]	-0.72 [-2.10]	0.58 [4.02]	0.06 [0.94]	No
DOLS	1.51 [6.59]	0.22 [1.83]	-0.40 [-0.49]	0.56 [2.42]	0.04 [0.41]	

Multivariate Cointegration Approach

	Trace	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	2	0	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	3	0	See below

Vector Error Correction Model[185]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.48 [-4.68]	1.65 [12.31]	0.35 [8.04]	-1.21 [-4.52]	0.77 [6.56]	0.06 [1.19]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.36 [-4.00]	2.66 [13.68]	0.44 [9.38]	0.38 [1.26]	3.13 [8.79]	0.27 [5.02]	Yes

[183] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[184] t-values are given in squared-brackets.
[185] The test regressions take no lagged-differences of model variables.

The ECM estimates an error correction coefficient of -0.59, indicating that over half of the total misalignment/departure of rer_def_nt from its long-run equilibrium is corrected by the next year. Thus, the necessary conditions are met for me to proceed with FMOLS and DOLS estimation of the long-run model. The two estimators produce similar results: a statistically insignificant BS effect, with $\tilde{\alpha}$ close to zero and having a t-value less than one. Consequently, there is no convincing empirical evidence in favour of a BS effect for Korea using the single equation framework.

An outcome somewhat more favourable to the BS hypothesis is obtained when I use the multivariate cointegration framework. Trace statistics associated with the Johansen ML test finds evidence of cointegration in both Case 3 and Case 4. This allows me to proceed with VEC estimation of the long-run model elasticities.

Specification 4 of the VEC model produces evidence to support the existence of a BS effect for Korea. The model qualifies because it fulfils both necessary conditions: (a) the EC coefficient is negative and statistically significant, and (b) the long-run BS coefficient is positive and highly significant. These results conflict with those for Case 3, where the EC coefficient meets the necessary condition, but the BS coefficient is statistically insignificant. Thus the multiple cointegration framework produces mixed results for Korea.

Turning now to the other determinants of long-run real exchange rates, I find strong evidence against the PPP model, consistent with earlier findings from Chapter Eight. The coefficient for rer_def_t is highly statistically significant in both Case 3 and Case 4 specifications. With respect to the demand-side determinants, there is strong statistical support in favour of real oil prices and real GDP per capita affecting the country's real exchange rate. In contrast, real government spending shows mixed evidence in regards to real exchange rate movements.

9.4.4 Malaysia

Malaysia was the other country (in addition to Indonesia) for which my previous analysis in Chapter Eight was unable to reject PPP. Accordingly, I begin by estimating the BS model which incorporates the PPP restrictions. The results from this analysis are reported in TABLE 9.6.

The single equation tests provide evidence for cointegration, though only for the ECM estimates. The EG residual test does not find evidence of cointegration. However, the EC coefficient in the respective single equation model is negative and statistically significant. As a result, I use FMOLS and DOLS to estimate the long-run relationships. The two estimating procedures reach the same conclusion. Both produce statistically significant, albeit wrong-signed (negative) estimates of the BS coefficient. As a result, the single equation results are unanimous in finding no evidence for BS.

There is also support in favour of long-run co-movement between rer_def_nt and its respective demand-side determinants when the model is estimated using the multivariate cointegration approach. Both the Trace and the Maximum Eigenvalue statistics of Cases 3 and 4 find evidence of either two or three cointegrating vectors. This allows me to proceed with estimating the VECM.

While there is evidence of cointegration, there is no evidence that rer_def_nt adjusts in response to deviations from long-run equilibrium. The EC coefficients in the rer_def_nt equation are statistically insignificant in both the Case 3 and Case 4 specification. Further the BS coefficients are negative and statistically significant. As a result, the multivariate cointegration results, like the single equation results before them, do not produce evidence of a BS effect for Malaysia -- at least when the associated model incorporates PPP.

TABLE 9.6: Cointegration Tests Results for Malaysia[186] (1980-2013)[187]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)***	I(1)
$\tilde{a}$	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[188]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.39 [-4.08]

Estimator	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	0.26 [3.60]	0.17 [0.58]	0.10 [0.76]	-0.44 [-6.65]	No
DOLS	0.15 [2.20]	0.51 [1.89]	0.10 [0.79]	-0.44 [-7.89]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	3	3	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	2	See below

Vector Error Correction Model[189]

$EC_{rer\_def\_nt}$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR					
-0.05 [-0.63]	-0.10 [-0.98]	1.58 [3.95]	0.35 [2.01]	-0.35 [-3.77]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR					
-0.04 [-1.30]	1.14 [4.90]	-4.89 [-4.73]	-6.41 [-4.13]	-0.48 [-2.13]	No

[186] The country does not include rer_def_t as a model regressor, since PPP holds for its traded sector prices against U.S. (see Chapter 8).
[187] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[188] t-values are given in squared-brackets.
[189] The test regressions take no lagged-differences of model variables.

The demand-side variables do not present a consistent picture in the estimates of TABLE 9.6. While the real price of oil (*rop*) is positive and significant in both single equation estimates, it is insignificant in one of the multivariate cointegration specifications. Government expenditures are positive and significant in the Case 3 VECM, but negative and significant in the Case 4 specification. And so on.

As discussed above, just because I was unable to reject PPP for Malaysia in my Chapter Eight analysis, that does not mean that the alternative hypothesis is not also a reasonable possibility. Accordingly, I repeat the preceding analysis using the modified version of the BS model, augmented with demand-side variables. These results are reported in TABLE 9.7.

Inclusion of rer_def_t as a long-run determinant of real exchange rates does not win any additional support for the BS hypothesis. This is true for both single equation and multivariate cointegration tests. FMOLS and DOLS produce negative slope coefficients for $\tilde{a}$, similar to the preceding analysis (see TABLE 9.7). VEC test findings are also no different from the earlier results (see TABLE 9.7), producing insignificant EC coefficients and significant but negative BS coefficients. Thus, under any of the respective model specifications, there is no evidence of a BS effect for Malaysia.

9.4.5 Pakistan

The single equation cointegration test results for Pakistan are given in TABLE 9.8. The tau-statistic, generated by the EG residual test, is unable to reject the null hypothesis of no cointegration. In contrast, the ECM does find evidence of cointegration among model variables. However, upon estimating the long-run BS effect, FMOLS and DOLS estimators produce negative coefficients for $\tilde{a}$, with the DOLS estimate being statistically significant. Thus, single equation cointegration methods do not support the BS hypothesis for the country.

TABLE 9.7: Cointegration Tests Results for Malaysia (1980-2013)[190]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
gdp_percapita	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[191]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.39 [-4.08]

Estimator	*rer_def_t*	*rop*	*gov_exp*	*gdp_percapita*	$\tilde{a}$	Does BS Effect Hold?
FMOLS	1.33 [4.29]	0.42 [6.22]	0.14 [0.63]	0.16 [1.59]	-0.26 [-4.08]	No
DOLS	1.60 [5.30]	0.38 [4.45]	0.71 [3.37]	0.50 [4.82]	-0.28 [-6.15]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	3	2	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	4	4	See below

Vector Error Correction Model[192]

$EC_{rer\_def\_nt}$	*rer_def_t*	*rop*	*gov_exp*	*gdp_percapita*	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.03 [-0.40]	0.49 [0.94]	0.00 [0.02]	1.58 [4.12]	0.29 [1.73]	-0.26 [-2.25]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
0.02 [1.03]	-3.01 [-1.33]	1.15 [2.44]	-8.13 [-5.08]	-11.94 [-4.26]	-0.72 [-1.79]	No

[190] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[191] t-values are given in squared-brackets.
[192] The test regressions take no lagged-differences of model variables.

TABLE 9.8: Cointegration Tests Results for Pakistan (1973-2008)[193]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)**	I(1)*	I(1)
rer_def_t	Yes	I(1)**	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[194]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.36 [-2.47]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	1.11 [7.98]	0.01 [0.35]	-0.46 [-7.88]	0.64 [9.83]	-0.04 [-0.58]	No
DOLS	1.41 [9.81]	0.14 [2.79]	-0.52 [-10.62]	0.50 [8.01]	-0.36 [-2.57]	

Multivariate Cointegration Approach

	Trace	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	1	1	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	2	2	See below

Vector Error Correction Model[195]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.19 [-1.41]	1.51 [8.30]	0.26 [5.28]	-0.41 [-5.57]	0.48 [5.63]	-0.34 [-3.36]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.17 [-1.30]	1.60 [8.59]	0.29 [5.48]	-0.30 [-2.03]	-0.09 [-0.17]	-0.35 [-3.39]	No

[193] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[194] t-values are given in squared-brackets.
[195] The test regressions take no lagged-differences of model variables.

There is sufficient support in favour of long-run co-movement between model variables using the multivariate method. The Trace and Maximum Eigenvalue statistics for both Case 3 and 4 of the Johansen ML test favour the existence of one or two cointegrating vectors. This allows me to estimate short-run to long-run dynamics of the model using the VECM.

In line with my previous country results, both cases of the VEC model do not support the existence of a causal relationship between rer_def_nt and its proposed determining factors. The test produces a statistically insignificant EC coefficient rer_def_nt equation. This implies that the underlying condition of the model is not met; i.e., short-lived fluctuations in rer_def_nt are not corrected by the series itself to make the series return to its long-run equilibrium. The long-run BS coefficient is also negative, displaying a counter-intuitive relationship between model variables. What's worse (for the BS hypothesis), it is also statistically significant. Thus, I conclude that there is no evidence in favour of the BS hypothesis for Pakistan.

With respect to the other regressors from the single equation and multivariate cointegration models, I find that rer_def_t is positively and significantly related to the trend behaviour of rer_def_nt. The three demand-side factors are also generally estimated to significantly affect real exchange rate movements in long-run. Government spending tends to induce appreciation in relative prices of nontradables whereas GDP per capita is inducing pressures of mixed nature (appreciation as well as depreciation) on rer_def_nt to make the series see trend departures from its long-run equilibrium levels.

9.4.6 Philippines

The empirical estimates for the Philippines are reported in TABLE 9.9 below. Overall, there are no traces of BS effect existing for the country, according to single equation cointegration test estimates.

TABLE 9.9: Cointegration Tests Results for Philippines (1971-2013)[196]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
gdp_percapita	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[197]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.10 [-1.09]

Estimator	*rer_def_t*	*rop*	*gov_exp*	*gdp_percapita*	$\tilde{a}$	Does BS Effect Hold?
FMOLS	-	-	-	-	-	No
DOLS	-	-	-	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	0	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace Statistics	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	0	See below

Vector Error Correction Model[198]

$EC_{rer\_def\_nt}$	*rer_def_t*	*rop*	*gov_exp*	*gdp_percapita*	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.00 [0.08]	-4.29 [-6.21]	-1.45 [-7.75]	-4.28 [-7.91]	1.71 [1.61]	-0.55 [-1.86]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
0.02 [0.23]	2.53 [15.67]	0.54 [10.92]	0.32 [2.52]	-1.49 [-5.85]	-0.19 [-2.10]	No

[196] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[197] t-values are given in squared-brackets.
[198] The test regressions take no lagged-differences of model variables.

The estimated tau-statistic of the model (in the EG test specification), when compared against the Mackinnon critical value, could not reject the null hypothesis of no cointegration. Similar lack of evidence for cointegration is given by the ECM. Thus, the single equation cointegration results argue against a BS effect for the Philippines.

The results obtained through the multivariate model are different from earlier ones. The Trace statistics for both specifications of the Johansen ML cointegration test (Case 3 and 4) identify two valid cointegrating vectors for the estimated models. This creates the opportunity for a BS effect to exist. Accordingly, I conduct VEC estimation.

The VEC model results are inconsistent with the theoretical predictions of the BS hypothesis. The model pre-condition is not satisfied. The rer_def_nt series does not show evidence of correcting short-run deviations from a long-run equilibrium. This nullifies the possibility of a BS effect for the Philippines. Further, the long-run BS coefficient is statistically significant but negative, suggesting a counter-intuitive association between $\tilde{a}$ and rer_def_nt for the Philippines and the U.S.

Turning to the role of inter-country tradables prices in explaining real exchange rate movements, appreciation in the rer_def_nt series is significantly driven by trend movements in rer_def_t, as is evident from the positive and significant coefficients for this variable in the single equation and multiple cointegration models. As for demand-side factors, shifts in rer_def_nt show evidence of being significantly associated with demand-side factors (though the coefficient for GDP per capita has contrasting signs depending on the specification).

9.4.7 Singapore

The country results for Singapore are reported in TABLE 9.10. Starting from the single equation cointegration tests, there is no evidence of a long-run association between rer_def_nt

and its respective determinants, with consistent results being reported by the both the EG residual test and estimates from the ECM. Thus, the single equation cointegration approach does not produce any evidence to support the BS hypothesis.

The rank test results from the multivariate cointegration approach present a somewhat different picture. Cases 3 and 4 of the Johansen ML test find evidence of two and one cointegrating vectors according to the Trace and Maximum Eigenvalue statistics, respectively. These results provide sufficient reason to estimate the VEC model in searching for evidence of a BS effect.

The VEC model rejects the BS hypothesis as the two model conditions are not sufficiently satisfied. In the Case 3 specification, trend movements of rer_def_nt do not adjust to short-run deviations so that long-run equilibrium can be re-established. This invalidates the error correction part of the model, necessary to establish a valid BS effect. In addition, the BS coefficient is estimated to be negative, suggesting that $\tilde{a}$ is affecting rer_def_nt in a manner inconsistent with the BS hypothesis. This is true for both specifications of the model.

The other model regressors are generally estimated to be statistically significant determinants of rer_def_nt. However, the associated estimates are sometimes inconsistent across model specifications. For example, the estimated coefficient for rer_def_t is negative and significant in the Case 3 specification, but positive and significant in the Case 4 specification. Government spending also shows conflicting coefficient signs across equations.

TABLE 9.10: Cointegration Tests Results for Singapore (1970-2006)[199]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(0)**	I(0)**	I(0)
rer_def_t	Yes	I(0)**	I(0)**	I(0)
$\tilde{a}$	Yes	I(1)**	I(0)**	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[200]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.09 [-0.96]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	-	-	-	-	-	No
DOLS	-	-	-	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	1	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	1	See below

Vector Error Correction Model[201]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.03 [2.67]	-20.78 [-4.96]	-0.45 [-0.94]	13.86 [4.18]	-9.43 [-5.46]	-4.69 [-7.70]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.49 [-3.37]	1.67 [6.11]	0.26 [8.04]	-1.72 [-6.42]	-4.20 [-12.41]	-0.26 [-6.81]	No

[199] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[200] t-values are given in squared-brackets.
[201] The test regressions take no lagged-differences of model variables.

9.4.8 Sri Lanka

For Sri Lanka, the EG and ECM cointegration test results, reported in TABLE 9.11, do not provide support for the presence of a BS effect for the country. There are no signs of significant long-run cointegration between rer_def_nt and its long-run determinants in the single equation cointegration estimates. In contrast, the multivariate cointegration findings do find evidence of cointegration. The Trace and Maximum Eigenvalue statistics of Cases 3 and 4 of the Johansen ML test estimate ranks of three and two, respectively, supporting the existence of multiple valid cointegrating vectors between model variables.

However, support for the BS hypothesis comes to an end with estimation of the VECM. One of the two necessary conditions for the BS hypothesis is not met. The EC coefficient in the rer_def_nt equation is statistically indistinct from zero in both specifications of the model, indicating that rer_def_nt does not self-correct deviations from long-run equilibrium. This invalidates the possibility of a BS effect.

Sri Lanka produces mixed evidence on the role of tradables prices, as evident from a comparison of the estimated coefficients for rer_def_t in the Case 3 and Case 4 specifcations of the VECM. Only specification 3 produces a statistically significant coefficient for rer_def_t. With respect to demand-side variables, real oil prices and government expenditures show consistent, statistically significant coefficient estimates across the two specifications, while the estimates for GDP per capital differ in sign and statistical significance in the two cases.

TABLE 9.11: Cointegration Tests Results for Sri Lanka (1981-2010)[202]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(0)***	I(0)***	I(0)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[203]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
rer_def_nt	No	1	-0.31 [-1.06]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	-	-	-	-	-	No
DOLS	-	-	-	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	3	2	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
rer_def_nt	2	2	See below

Vector Error Correction Model[204]

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
0.00 [0.78]	13.06 [3.04]	5.62 [6.42]	-11.70 [-2.34]	2.61 [0.85]	13.13 [4.77]	No

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
0.01 [1.14]	5.12 [1.53]	4.46 [6.30]	-8.19 [-4.41]	-18.89 [-1.76]	0.66 [1.93]	No

202 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
203 t-values are given in squared-brackets.
204 The test regressions take no lagged-differences of model variables.

9.4.9 Thailand

The empirical evidence on the BS effect for Thailand lines up behind the results I have obtain for my other sample countries. The results are reported in TABLE 9.12. The Engle-Granger (EG) residuals show evidence of being a unit root process. Nevertheless, evidence in favour of cointegration is provided by the ECM. Consequently, I use FMOLS and DOLS to estimate the long-run relationship between rer_def_nt and the other model variables. Both estimators produce significant but negative estimates for the BS coefficient. Thus, single equation cointegration estimation produce no support for the BS hypothesis for Indonesia.

In contrast to the single equation models, the multivariate cointegration model produces tidy results in favour of cointegration. Both Trace and Maximum Eigenvalue tests identify a single cointegrating equation for both Case 3 and Case 4 specifications. Further, the associated EC terms in the rer_def_nt equations show evidence that this series adjusts to deviations from long-run equilibrium as is required for the BS hypothesis to hold. Nevertheless, the estimated coefficient on the productivity differential variable, $\tilde{a}$, is wrong-signed and highly significant. This provides evidence against the BS hypothesis. Thus, across both single equation and multivariate cointegration approaches, there is no evidence that the BS hypothesis holds for Thailand.

Turning to the other model regressors, I find consistent evidence that rer_def_t is positively and significantly related to rer_def_nt. Real GDP per capita is also a positive and significant determinant of rer_def_nt, while estimates for the other variables vary in sign and significance across estimation methods and model specifications.

TABLE 9.12: Cointegration Tests Results for Thailand (1971-2013)[205]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
rer_def_nt	Yes	I(1)***	I(1)***	I(1)
rer_def_t	Yes	I(1)***	I(1)***	I(1)
$\tilde{a}$	Yes	I(1)***	Greater than I(1)	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)**	I(1)**	I(1)
$gdp\_percapita$	Yes	I(1)**	I(1)***	I(1)

Single Equation Cointegration Approach[206]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient[207]
rer_def_nt	No	1	-0.33 [-1.93]

Estimator	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
FMOLS	1.31 [5.99]	0.07 [1.24]	0.15 [8.16]	0.72 [5.51]	-0.18 [-2.69]	No
DOLS	1.30 [4.37]	0.15 [1.20]	-0.10 [-0.25]	0.74 [3.37]	-0.19 [-1.60]	

Multivariate Cointegration Approach

	Trace	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
rer_def_nt	1	1	See Below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
rer_def_nt	1	1	See below

Vector Error Correction Model[208]

$EC_{rer\_def\_nt}$	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a}$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.14 [-1.96]	1.09 [3.78]	0.24 [3.47]	-0.52 [-2.31]	0.48 [2.82]	-0.46 [-4.81]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.10 [-2.44]	0.82 [1.75]	0.30 [2.66]	0.25 [0.54]	1.27 [2.27]	-0.93 [-5.89]	No

205 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
206 t-values are given in squared-brackets.
207 The test regression also include first lagged-difference of model regressors.
208 The test regressions take no lagged-differences of model variables.

9.4.10 Summary of Individual Country Studies

In the empirical literature, demand-side factors have received much attention as determinants of the trend behaviours of real exchange rates. In this chapter, I estimated an augmented version of the BS hypothesis by incorporating some widely recognized demand-side determinants of real exchange rates into a modified version of the BS model (introduced in Chapter Eight). In light of the lack of evidence in support of the BS hypothesis from previous investigations, I explored whether the incorporation of demand-side factors into the model could unearth evidence of a BS effect. Previous research has found that supply-side effects can be deficient in explaining real exchange rate movements unless combined with demand-side determinants of home and foreign prices (De Gregorio and Wolf, 1994; Chinn and Johnston, 1996).

However, this line of inquiry did not produce appreciably more support for the BS hypothesis. Korea is the only country in my sample whose real exchange rates against the U.S. showed evidence consistent with the implications of the BS theory, but even this evidence was mixed. Arguably, partial support for the BS hypothesis is provided by Indonesia, but only when the estimated models allow for failure of PPP, despite the fact that the null hypothesis of PPP cannot be rejected for Indonesia.

On the whole, the findings from this chapter are comparable with my results from preceding chapters. I note that my results for Korea are in line with empirical evidence from earlier studies (Chinn, 1996, 2000, Thomas and King, 2008). Further, my inability to find strong support in favor of the BS hypothesis for Asia (except Japan and Korea) generally matches the findings of Ito et al. (1999) and Wang et al. (2016), who also investigated Asia-Pacific countries for a BS effect.

My analysis was relatively more successful in identifying a role for demand-side variables, as well as the inter-country tradables price ratio, as determinants of long-run trend behaviour in real exchange rates. Multivariate cointegration tests frequently identified long-run associations between demand-side shocks and the trend movements of rer_def_nt. There was also consistent evidence across the majority of sample countries against the validity of PPP between inter-country traded sector prices.

As in previous chapters, I will proceed with an investigation of Hong Kong, and then go on to pool my data in a final attempt to find evidence for the BS hypothesis.

TABLE 9.13: Does Augmented Version of the Balassa-Samuelson Hypothesis Hold Valid? Summary of Results by Country

Country	Individual Results: *Single Equation Cointegration Method*	Individual Results: *Multivariate Cointegration Method* *Case 3*	Individual Results: *Multivariate Cointegration Method* *Case 4*	Country Summary
Indonesia (1976-2013)	No	No	No	No
Indonesia_PPP[209] (1976-2013)	Mixed	No	No	Mixed
Japan (1970-2013)	No	No	No	No
Korea (1970-2013)	No	No	Yes	Mixed
Malaysia (1980-2013)	No	No	No	No
Malaysia_PPP[210] (1980-2013)	No	No	No	No
Pakistan (1973-2008)	No	No	No	No
Philippines (1971-2013)	No	No	No	No

[209] The country analysis includes rer_def_t as model regressor to control for plausible absence of PPP for inter-country tradables prices.
[210] The country analysis includes rer_def_t as model regressor to control for plausible absence of PPP for inter-country tradables prices.

Country	**Individual Results**			**Country Summary**
	Single Equation Cointegration Method	*Multivariate Cointegration Method*		
		Case 3	*Case 4*	
Singapore (1970-2006)	No	No	No	No
Sri Lanka (1981-2010)	No	No	No	No
Thailand (1971-2013)	No	No	No	No

9.4.11 Hong Kong

Similar to proceeding chapters, the country study of Hong Kong is conducted separately. The reason lies with the dissimilar scheme of sectoral divisions of the country, which does not follow from Dumrongrittikul's (2012) study, and so must be investigated independently.

Starting with the first sectoral classification type (HKG_1), and employing single equation cointegration tests, I find no evidence of a long-run association between $rer\_def\_nt\_1$ and $\tilde{a}\_1$. The tau statistic for the EG residuals test does not allow one to reject the null hypothesis of no cointegration with appropriate statistical significance. A similar conclusion follows from the ECM, which produces a statistically insignificant EC coefficient, so that one cannot reject the null hypothesis of no cointegration. Thus, I conclude there is no BS effect for Hong Kong when estimating single equation cointegration models and using the first sectoral division scheme.

Evidence in favour of cointegration is provided by the multivariate cointegration models. Both specifications of the Johansen ML test reveal the existence of valid cointegrating vectors, as suggested by both Trace and Maximum Eigenvalue statistics. This allows me to estimate VEC models in an attempt to detect the presence of a BS effect for the country.

The two specifications of the VEC model (Cases 3 and 4) produce similar results. While both specifications find evidence of the necessary error correction behaviour, the estimated BS coefficient is inconsistent with the BS hypothesis. In the Case 3 specification, the estimated coefficient for $\tilde{a}\_1$ is statistically insignificant. In the Case 4 specification, it is significant but negative. Overall, I conclude that there is a lack of evidence in support of the BS hypothesis for Hong Kong when using the first type of sectoral division.

TABLE 9.14: Cointegration Tests Results for Hong Kong: Case 1 (1980-2008)[211]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_1$	Yes	Greater than I(1)	I(1)**	Inconclusive
$rer\_def\_t\_1$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_1$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[212]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
$rer\_def\_nt\_1$	No	1	-0.08 [-0.48]

Estimator	$rer\_def\_t\_1$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_1$	Does BS Effect Hold?
FMOLS	-	-	-	-	-	No
DOLS	-	-	-	-	-	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_1$	4	2	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_1$	4	3	See below

Vector Error Correction Model[213]

$EC_{rer\_def\_nt\_1}$	$rer\_def\_t\_1$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_1$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
-0.22 [-1.99]	3.13 [9.67]	-0.01 [-0.41]	1.78 [8.81]	1.86 [15.10]	0.01 [0.07]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.21 [-2.56]	2.40 [5.16]	0.11 [1.61]	3.71 [7.03]	3.07 [6.04]	-0.40 [-2.25]	No

211 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
212 t-values are given in squared-brackets.
213 The test regression takes zero lagged-differences of model variables.

The two specifications of the VEC model yield long-run coefficients for $rer\_def\_t\_1$ that are significantly different from zero. As a result, consistent with my earlier analysis for HKG_1, I reject the PPP hypothesis for Hong Kong. This is evidence against the classical formulation of BS theory, which assumes no role for tradables prices in displacing real exchange rates from their long-run equilibrium. With respect to the demand-side determinants of real exchange rates, I find that both gov_exp and $gdp\_percapita$ are consistently estimated to be positive and significant determinants of $rer\_def\_nt\_1$.

I next discuss the empirical estimates for the other three sectoral classifications of Hong Kong. However, I will mostly discuss the results collectively, as the three classifications produce generally similar results. The EG single equation cointegration test, under all three sectoral classifications of Hong Kong, concludes that rer_def_nt is cointegrated with the other model variables. However, when I use FMOLS and DOLS to estimate the long-run relationships, the long-run slope coefficient fir $\tilde{\alpha}$ is always negative and/or statistically insignificant. Thus, for all three types of sectoral divisions, there is no evidence of a BS effect when using single equation cointegration methods.

In contrast, the results from the multivariate cointegration analysis produce evidence that supports the existence of cointegration between model variables. This is true for all three sectoral classifications. The two types of test statistics (Trace and Maximum Eigenvalue statistics) unanimously support the existence of multiple cointegrating vectors, under both Case 3 and Case 4 of the Johansen ML test. These findings allow me to proceed with VEC estimation as a next step.

In every case but one (the exception being Case 4, TABLE 9.17), the VECM estimates indicate an invalid EC process as the associated EC coefficient is positive and significant. Furthermore, the estimated BS coefficients are negative and statistically significant, completely counter to the BS hypothesis. The only exception is Case 4 of HKG_4, where the EC term and the BS coefficient are consistent with the presence of a BS effect. Taken together, the majority of evidence argues against the BS hypothesis.

Regarding the effect of inter-country traded sector prices on real exchange rate movements, there is again sizable statistical support that the long-run coefficient of rer_def_t is positive and significantly related to rer_def_nt. Somewhat curiously, the only exception is in the specification that supports the BS hypothesis (Case 4, TABLE 9.17). The estimates for demand-side factors, while often statistically significant, sometimes vary in sign across specification and estimation method.

TABLE 9.15: Cointegration Tests Results for Hong Kong: Case 2 (1980-2008)[214]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_2$	Yes	I(1)**	I(1)**	I(1)
$rer\_def\_t\_2$	Yes	I(1)**	I(1)**	I(1)
$\tilde{a}\_2$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[215]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
$rer\_def\_nt\_2$	Yes	1	-0.30 [-1.39]

Estimator	$rer\_def\_t\_2$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_2$	Does BS Effect Hold?
FMOLS	0.74 [7.53]	0.11 [7.77]	0.56 [10.38]	0.43 [12.51]	0.03 [0.65]	No
DOLS	0.91 [1.87]	0.07 [0.77]	0.34 [0.98]	0.76 [1.93]	0.01 [0.05]	

Multivariate Cointegration Approach

	Trace	Max Eigenvalue	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR			
$rer\_def\_nt\_2$	4	3	See below
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR			
$rer\_def\_nt\_2$	4	3	See below

Vector Error Correction Model[216]

$EC_{rer\_def\_nt\_2}$	$rer\_def\_t\_2$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_2$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.41 [6.42]	1.66 [12.23]	-0.01 [-0.74]	0.57 [7.57]	0.77 [14.64]	-0.16 [-2.61]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
0.02 [5.14]	17.43 [8.15]	0.20 [0.49]	16.98 [7.18]	22.74 [6.56]	-3.33 [-3.34]	No

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214 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
215 t-values are given in squared-brackets.
216 The test regression takes zero lagged-differences of model variables.

TABLE 9.16: Cointegration Tests Results for Hong Kong: Case 3 (1980-2008)[217]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_3$	Yes	I(1)**	I(1)***	I(1)
$rer\_def\_t\_3$	Yes	I(1)**	I(0)**	Inconclusive
$\tilde{a}\_3$	Yes	I(1)**	I(0)**	Inconclusive
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)***	I(1)***	I(1)
$gdp\_percapita$	Yes	I(1)***	I(1)***	I(1)

Single Equation Cointegration Approach[218]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient
$rer\_def\_nt\_3$	Yes	1	-0.35 [-1.51]

Estimator	$rer\_def\_t\_3$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_3$	Does BS Effect Hold?
FMOLS	1.32 [12.61]	0.04 [1.26]	0.46 [4.15]	0.18 [0.90]	-0.07 [-0.91]	No
DOLS	1.20 [7.24]	0.09 [1.33]	0.62 [2.93]	0.31 [0.57]	-0.07 [-0.47]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_3$	5	3	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_3$	5	2	See below

Vector Error Correction Model[219]

$EC_{rer\_def\_nt\_3}$	$rer\_def\_t\_3$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_3$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.45 [5.87]	1.53 [12.91]	-0.04 [-0.88]	0.21 [1.74]	1.10 [4.56]	-0.27 [-2.85]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
0.19 [1.80]	0.57 [4.74]	-0.00 [-0.11]	-1.18 [-5.89]	-0.16 [-4.84]	-0.45 [-4.53]	No

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217 *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
218 t-values are given in squared-brackets.
219 The test regression takes zero lagged-differences of model variables.

TABLE 9.17: Cointegration Tests Results for Hong Kong: Case 4 (1980-2008)[220]

ADF and DF-GLS Unit Root Tests

Variables	White Noise Residuals	ADF Test	DF-GLS Test	Conclusion
$rer\_def\_nt\_4$	Yes	Greater than I(1)	I(0)**	Inconclusive
$rer\_def\_t\_4$	Yes	Greater than I(1)	Greater than I(1)	Greater than I(1)
$\tilde{a}\_4$	Yes	I(1)***	I(1)***	I(1)
rop	Yes	I(1)***	I(1)***	I(1)
gov_exp	Yes	I(1)**	I(1)**	I(1)
$gdp\_percapita$	Yes	I(1)**	I(1)***	I(1)

Single Equation Cointegration Approach[221]

Dependent Variable	Are EG Test Residuals I(0)?	Lags of regressand	EC Coefficient[222]
$rer\_def\_nt\_4$	Yes	1	-0.12 [-0.71]

Estimator	$rer\_def\_t\_4$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_4$	Does BS Effect Hold?
FMOLS	0.95 [6.46]	0.28 [13.10]	0.89 [10.33]	1.32 [20.15]	-0.11 [-1.99]	No
DOLS	1.31 [4.71]	0.14 [3.30]	0.78 [4.91]	1.93 [9.05]	-0.07 [-0.61]	

Multivariate Cointegration Approach

Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_4$	4	2	See below

Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR

	Trace	Max Eigenvalue	Does BS Effect Hold?
$rer\_def\_nt\_4$	5	3	See below

Vector Error Correction Model[223]

$EC_{rer\_def\_nt\_4}$	$rer\_def\_t\_4$	rop	gov_exp	$gdp\_percapita$	$\tilde{a}\_4$	Does BS Effect Hold?
Case 3: Linear deterministic trend in the data and an intercept in CE and test VAR						
0.11 [5.27]	5.54 [15.17]	-0.13 [-2.23]	0.84 [3.85]	2.02 [11.25]	-1.44 [-8.67]	No
Case 4: Linear deterministic trend in data, intercept and trend in CE and no trend in VAR						
-0.01 [-4.37]	-6.86 [-12.44]	21.69 [2.65]	-14.01 [-3.29]	-24.33 [-3.79]	3.15 [7.76]	Yes

[220] *, ** and *** are showing significance of coefficients at 10%, 5% and 1% significance level respectively.
[221] t-values are given in squared-brackets.
[222] The test regression also include first lagged-difference of model regressors.
[223] The test regression takes zero lagged-differences of model variables.

9.5 Panel Data Estimation Results

I start my panel data estimations by formally testing the respective model variables for unit roots. As always, I am testing the variables using the Fisher-ADF and Fisher-PP unit root tests, and the Hadri stationarity test.

The results are reported in the first panel of TABLE 9.18. For rer_def_nt, the variable is unit root in levels according to the Fisher-PP and the Hadri stationarity test. The results for the Fisher-ADF test do not support this conclusion, as the test indicates that the variable is integrated of order zero in levels. However, I conclude that the sum of the evidence indicates that the series is a unit root process in levels.

For rer_def_t, all three tests unanimously agree that the series is a unit root process in levels. For $\tilde{a}$, the two unit root tests indicate that the series is level stationary, while the Hadri stationarity test suggests that the order of integration is greater than 1. Again, I conclude that the weight of the evidence indicates that the variable is stationary in levels. With respect to rop, gov_exp and $gdp\_percapita$, while the three tests do not always agree, the weight of the evidence points to all three being unit root in levels.

The second and third panels of TABLE 9.18 display the test results for the Pedroni residual based cointegration test and the Johansen-Fisher panel cointegration test, respectively. Discussing the Pedroni test results first, I opted for automatic lag selection. None of the seven test statistics provide evidence in support of cointegration between model variables.

TABLE 9.18: Summary of Results for Panel Unit Root and Cointegration Tests Augmented Version of the Balassa-Samuelson Hypothesis[224,225,226]

Panel Unit Root Test Results (Order of Integration as Determined by)

Variables	Fisher-ADF	Fisher-PP	Hadri	Conclusion
rer_def_nt	I (0)*	I (1)***	I (1)***	I (1)
rer_def_t	I (1)***	I (1)***	I (1)***	I(1)
$\tilde{a}$	I (0)**	I (0)**	Greater than I (1)	I(0)
rop	I (1)***	I (1)***	Greater than I(1)	I(1)
gov_exp	I (1)***	I (1)***	Greater than I(1)	I(1)
$gdp\_percapita$	I (1)***	I (1)***	I(1)***	I(1)

Pedroni Panel Cointegration Test Results[227]

Common AR Coefficients (Within Dimension)				*Individual AR Coefficients (Between Dimension)*			
Panel v Statistics	Panel ρ Statistics	Panel PP Statistics	Panel ADF Statistics	Group ρ Statistics	Group PP Statistics	Group ADF Statistics	Does BS Effect Hold?
1.12	2.06	0.76	0.52	2.83	1.03	1.09	No

224 ***, ** and * are representing significance of sample statistics at 1%, 5% and 10% levels respectively.
225 Hong Kong is omitted from panel estimations.
226 Indonesia and Malaysia are included in the panel for their models estimated in TABLE 9.2 and 9.5.2, respectively.
227 Pedroni panel cointegration is a test for null of no cointegration in both homogenous and heterogeneous panels. The test statistics are standardized and asymptotically normally distributed. See Pedroni (1995, 1999) for further details.

Johansen-Fisher Panel Cointegration Test Results[228,229]

	Case 3: Intercept (no trend) in cointegrating equation and VAR	
Fisher Stat (Trace)	Fisher Stat (Max Eigenvalue)	Does BS Effect Hold?
3	3	See below
	Case 4: Intercept and trend in cointegrating equation-no trend in VAR	
4	4	See below

Results for Panel FMOLS (PFMOLS) and Panel DOLS (PDOLS)[230] Estimators
Long-run Cointegrating Vectors for Augmented Version of the Balassa-Samuelson Hypothesis

Estimator			**Long-run Coefficient**[231]			**Does BS Effect Hold?**
	rer_def_t	rop	gov_exp	$gdp\_percapita$	$\tilde{a} = BS\ coefficient$	
PFMOLS	1.16 [9.10]	0.08 [1.98]	-0.21 [-1.74]	0.59 [8.94]	0.02 [0.60]	No
PDOLS	1.41 [11.58]	0.07 [1.78]	0.08 [0.57]	0.84 [12.21]	0.12 [2.83]	Yes

228 The test is maximum likelihood based rank test.
229 Lag selection is done through SIC under panel VAR.
230 Lead = Lag = 1,
231 t-values are given in squared-brackets.

In contrast, the Fisher-Johansen panel cointegration tests support the existence of a long-run association between model variables. As the test requires the user to specify lag length, I used SIC to determine the appropriate number of lags to put in. The Trace and Maximum Eigenvalue statistics for Cases 3 and 4 indicated 3 and 4 cointegrating vectors, respectively.

Because the Johansen-Fisher test results suggest cointegration, I proceed by estimating the long-run equilibrium relationship using panel FMOLS (PFMOLS) and panel DOLS (PDOLS). The two panel estimators produce dissimilar results. The PFMOLS estimator yields a statistically insignificant BS coefficient. In contrast, the PDOLS estimator produces a positive and statistically significant coefficient. This constitutes mixed support for the BS hypothesis. This finding is in contrast with my pooled data estimates from preceding chapters, which found no support for the BS hypothesis.

Finally, looking into the explanatory power of rer_def_t in determining the trend behaviour of rer_def_nt, the long-run elasticities of rer_def_t produced by the PFMOLS and PDOLS estimators indicate that the assumption of PPP is not valid for this panel of Asian countries. These findings are consistent with my previous results. Amongst demand-side factors, GDP per capita is most significant, and is estimated to positively contribute to rer_def_nt. The estimates for real oil prices are also positive and significant, but less so; while the PFMOLS estimates for government expenditures are negative and weakly significant, while the PDOLS estimates are insignificant.

9.5.1 Summary of Panel Data Results

This chapter tested an augmented version of the modified BS model, controlling for a number of demand-side variables that feature prominently in the literature. The results of this exercise produced somewhat greater support for the BS hypothesis. However, the resulting evidence is mixed, at best.

With somewhat greater confidence I conclude that the assumption of tradables PPP, which is central to the classical formulation of the BS hypothesis, is empirically rejected by the data, as the variable rer_def_t is consistently found to be a significant determinant of real exchange rates. Likewise, demand-side determinants are generally found to be statistically significant, though signs and significances are found to vary across estimation procedures and model specifications.

APPENDIX-E

EViews Programming Code for Korea

```
create y 1970 2013
'importing data from Excel for Korea
import  "C:\Users\Maryam\Desktop\BS Studies\PhD Thesis-II\EViews and STATA Progarm Codes\Chapter-9\Chapter 9.xlsx" range="Korea"

'*********************************************************************************************************************************
'ESTIMATING BALASSA-SAMUELSON EFFECT FOR rer_def_NT, rer_def_T, rop, gdppc, gexp & a_tilde
'*********************************************************************************************************************************
'*****************************************************************
'STEP 0: Tests for Unit Root in Individual Time Series
'*****************************************************************

'***************************
'Graph for Korea's rop
'***************************

genr rop = rop
freeze(figure_rop) rop.line
figure_rop.addtext(t) rop (Korea):  1970-2013
figure_rop.addtext(b) Year
figure_rop.addtext(l) rop
figure_rop.legend(off)

'We see from the FIGURE that rop has a time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'******************************************
'ADF Unit Root Test for Korea's rop
'******************************************

freeze(table_9_4_1_rop_adf) rop.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -2.11 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the adf test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,rop_adf) rop.uroot(adf,const,trend,info=sic)
freeze(rop_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the rop series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr ropdiff = d(rop)
freeze(figure_ropdiff) ropdiff.line
figure_ropdiff.addtext(t) Drop (Korea):  1970-2013
figure_ropdiff.addtext(b) Year
figure_ropdiff.addtext(l) Drop
figure_ropdiff.legend(off)

'The graph is not particularly illuminating.  Depending on how you look at it, it could have a time trend to it.  However, from the graphical inspection of the series, I conclude that it does not have a time trend, we run the ADF test with a constant only.
```

```
'So we begin the whole process over again:

genr ropdiff = d(rop)
freeze(table_9_4_2_ropdiff1_adf) ropdiff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0.  The unit root test produces a t-value of -7.01 which is now smaller than our 5% criterion -2.93.  Thus, we may now reject the null of nonstationarity in first differenced series of rop.  There is no reason to go further.  The last thing we do is check for white noise.

genr resid = 0
freeze(mode=overwrite,ropdiff1_adf) ropdiff.uroot(adf,const,info=sic)
freeze(ropdiff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the rop series is I(1).

'**********************************************
'DF-GLS Unit Root Test for Korea's rop
'**********************************************

freeze(table_9_4_3_rop_dfgls) rop.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -1.85 which is smaller than our 5% criterion -3.19.  Thus, at this point, we may reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr ropdiff = d(rop)
freeze(table_9_4_4_ropdiff1_dfgls) ropdiff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 0.  The unit root test produces a t-value of -7.00 which is now smaller than our 5% criterion -1.95.  Thus, we may not reject the null of nonstationarity in first differenced series of rop.

''Putting it all together, I conclude that the rop series is I(1), a finding compatible with my ADF test results.

'****************************
'Graph for Korea's  gexp
'*****************************

genr gexp = gexp
freeze(figure_gexp) gexp.line
figure_gexp.addtext(t) gexp (Korea):  1970-2013
figure_gexp.addtext(b) Year
figure_gexp.addtext(l) gexp
figure_gexp.legend(off)

'We see from the FIGURE that gexp has a time trend to it.  So we would include both an intercept and a time trend in our unit root regression equations.

'********************************************
'ADF Unit Root Test for Korea's gexp
'********************************************

freeze(table_9_4_1_gexp_adf) gexp.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -2.54 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the adf test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,gexp_adf) gexp.uroot(adf,const,trend,info=sic)
freeze(gexp_adf_correl) resid.correl
```

'Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the gexp series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test. I once again begin by graphing the (differenced) series.

```
genr gexpdiff = d(gexp)
freeze(figure_gexpdiff) gexpdiff.line
figure_gexpdiff.addtext(t) Dgexp (Korea):  1970-2013
figure_gexpdiff.addtext(b) Year
figure_gexpdiff.addtext(l) Dgexp
figure_gexpdiff.legend(off)
```

'The graph is not particularly illuminating. Depending on how you look at it, it could have a time trend to it. However, from the graphical inspection of the series, I conclude that it does not have a time trend, we run the ADF test with a constant only.

'So we begin the whole process over again:

```
genr gexpdiff = d(gexp)
freeze(table_9_4_2_gexpdiff1_adf) gexpdiff.uroot(adf,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p =0. The unit root test produces a t-value of -6.78 which is now smaller than our 5% criterion -2.93. Thus, we may now reject the null of nonstationarity in first differenced series of gexp. There is no reason to go further. The last thing we do is check for white noise.

```
genr resid = 0
freeze(mode=overwrite,gexpdiff1_adf) gexpdiff.uroot(adf,const,info=sic)
freeze(gexpdiff1_adf_correl) resid.correl
```

''Based on the Q-statistic, I conclude that the residuals are white noise. Putting it all together, I conclude that the gexp series is I(1).

'***
'DF-GLS Unit Root Test for Korea's gexp
'***

```
freeze(table_9_4_3_gexp_dfgls) gexp.uroot(dfgls,trend,info=sic)
```

'Note that the SIC automatic lag selection picks lags, p = 0. The unit root test produces a t-value of -2.63 which is smaller than our 5% criterion -3.19. Thus, at this point, we may reject the null of a unit root.

'Now let's see if the series is difference stationary or not

```
genr gexpdiff = d(gexp)
freeze(table_9_4_4_gexpdiff1_dfgls) gexpdiff.uroot(dfgls,const,info=sic)
```

'Note that the SIC automatic lag selection picks no lags, p = 0. The unit root test produces a t-value of -6.77 which is now smaller than our 5% criterion -1.95. Thus, we may not reject the null of nonstationarity in first differenced series of gexp.

''Putting it all together, I conclude that the gexp series is I(1), a finding compatible with my ADF test results.

'*******************************
'Graph for Korea's gdppc
'*******************************

```
genr gdppc = gdppc
freeze(figure_gdppc) gdppc.line
figure_gdppc.addtext(t) gdppc (Korea):  1970-2013
figure_gdppc.addtext(b) Year
figure_gdppc.addtext(l) gdppc
figure_gdppc.legend(off)
```

'We see from the FIGURE that gdppc has a time trend to it. So we would include both an intercept and a time trend in our unit root regression equations.

```
'*********************************************
'ADF Unit Root Test for Korea's gdppc
'**********************************************

freeze(table_9_4_1_gdppc_adf) gdppc.uroot(adf,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -0.10 which is greater than our 5% criterion -3.52.  Thus, at this point, we cannot reject the null of a unit root.

'Now, let's check for white noise.  To do that, I first set all the residuals = 0, then run the ADF test and finally will check for white noise.

genr resid = 0
freeze(mode=overwrite,gdppc_adf) gdppc.uroot(adf,const,trend,info=sic)
freeze(gdppc_adf_correl) resid.correl

'Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the gdppc series is not level stationary.

'The next thing I do is test whether the differenced series is stationary using the ADF test.  I once again begin by graphing the (differenced) series.

genr gdppcdiff = d(gdppc)
freeze(figure_gdppcdiff) gdppcdiff.line
figure_gdppcdiff.addtext(t) Dgdppc (Korea):  1970-2013
figure_gdppcdiff.addtext(b) Year
figure_gdppcdiff.addtext(l) Dgdppc
figure_gdppcdiff.legend(off)

'The graph is not particularly illuminating.  Depending on how you look at it, it could have a time trend to it.  However, from the graphical inspection of the series, I conclude that it does not have a time trend, we run the ADF test with a constant only.

'So we begin the whole process over again:

genr gdppcdiff = d(gdppc)
freeze(table_9_4_2_gdppcdiff1_adf) gdppcdiff.uroot(adf,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p =0.  The unit root test produces a t-value of -5.13 which is now smaller than our 5% criterion -2.93.  Thus, we may now reject the null of nonstationarity in first differenced series of gdppc.  There is no reason to go further.  The last thing we do is check for white noise.

genr resid = 0
freeze(mode=overwrite,gdppcdiff1_adf) gdppcdiff.uroot(adf,const,info=sic)
freeze(gdppcdiff1_adf_correl) resid.correl

''Based on the Q-statistic, I conclude that the residuals are white noise.  Putting it all together, I conclude that the gdppc series is I(1).

'*************************************************
'DF-GLS Unit Root Test for Korea's gdppc
'*************************************************

freeze(table_9_4_3_gdppc_dfgls) gdppc.uroot(dfgls,trend,info=sic)

'Note that the SIC automatic lag selection picks lags, p = 0.  The unit root test produces a t-value of -0.29 which is smaller than our 5% criterion -3.19.  Thus, at this point, we may reject the null of a unit root.

'Now let's see if the series is difference stationary or not

genr gdppcdiff = d(gdppc)
freeze(table_9_4_4_gdppcdiff1_dfgls) gdppcdiff.uroot(dfgls,const,info=sic)

'Note that the SIC automatic lag selection picks no lags, p = 0.  The unit root test produces a t-value of -5.04 which is now smaller than our 5% criterion -1.95.  Thus, we may not reject the null of nonstationarity in first differenced series of gdppc.
```

```
''Putting it all together, I conclude that the gdppc series is I(1), a finding compatible with my ADF test results.

'**********************************************
'Single Equation Cointegration Methods
'**********************************************

'**********************************************************
''Graph the suspected cointegrated series together
'**********************************************************
'The first step is to print out a graph of the series.  This is very important!

group g1 rer_def_NT rer_def_T rop gdppc gexp a_tilde
freeze(figure1) g1.line(x)
figure1.setelem(1) lcolor(black) symbol(1) lpat(1)
figure1.setelem(2) lcolor(black) symbol(4) lpat(1)
figure1.setelem(3) lcolor(black) symbol(7) lpat(1)
figure1.setelem(3) lcolor(black)
figure1.options linepat
figure1.addtext(t) rer_def_NT, lnpT & a_tilde (Korea & U.S): 1970-2013
figure1.addtext(b) Year
figure1.addtext(l) rer_def_NT
figure1.addtext(l) lnpT
figure1.addtext(r) a_tilde

''*******************************************************
''S1.A.Engle-Granger Approach to Cointegration
'*******************************************************

genr resid = 0
equation eg.ls rer_def_NT c rer_def_T rop gdppc gexp a_tilde
genr EC1 = resid

'First we test if the residuals of above regression are level stationary or not. If yes, next we'll proceed towards estimation of error correction model.

'***************************
'Graph for Korea's EC
'***************************
genr EC1 = EC1
freeze(figure_EC1) EC1.line
figure_EC1.addtext(t) EC1 (Korea):  1970-2013
figure_EC1.addtext(b) Year
figure_EC1.addtext(l) EC1
figure_EC1.legend(off)

'We see from the FIGURE that EC has time trend to it.  So we would include both an intercept and trend in our unit root regression equations.

'*******************************
'EG Test for Cointegration
'*******************************

freeze(table_9_4_EGC) g1.coint(method=eg)

'The null hypothesis will not be rejected as suggested by sample statistics.

''*******************************************
''S1.B.Error Correction Model (ECM)
'*******************************************
''**********************************************
'Selecting the number of lags in the VAR
'**********************************************
'NOTE: We do this because we need to have the "right" number of lags when it comes time to estimate our VEC model and test for cointegration.

var var1.ls 1 4   g1
freeze(var1_lagtest1) var1.laglen(4)
```

```
freeze(var1_lagtest2) var1.testlags

'The lag length test above indicates that the VAR has 1 lag.

var var2.ls 1 1  g1
freeze(var2_artest1) var2.correl
freeze(var2_artest2) var2.qstats(12)
freeze(var2_artest3) var2.arlm(12)

'The residuals are not white noise. But I can't go beyond this number of lags.

'We now try different lags of d(rer_def_T) and d(a_tilde), comparing SIC values across specifications.

genr resid = 0
equation eg.ls rer_def_NT c rer_def_T rop gdppc gexp a_tilde
genr ec1 = resid

var table_9_4_eg2a.ls 0 0 d(rer_def_NT)   @  c ec1(-1) d(rer_def_NT(-1))

var table_9_4_eg2b.ls 0 0 d(rer_def_NT)   @  c ec1(-1)  d(rer_def_NT(-1)) d(rer_def_T(-1)) d(rop(-1)) d(gdppc(-1)) d(gexp(-1)) d(a_tilde(-1))

'The evidence suggests that Model A is best.  Now we test that model for serial correlation.

var table_9_4_eg2a.ls 0 0 d(rer_def_NT)   @   c ec1(-1) d(rer_def_NT(-1))
freeze(table_9_4_eg2a_artest1) table_9_4_eg2a.correl
freeze(table_9_4_eg2a_artest2) table_9_4_eg2a.qstats(12)
freeze(table_9_4_eg2a_artest3) table_9_4_eg2a.arlm(12)

'The residuals are absolutely white noise.

''*************************
''Estimating EC Model
'**************************

'We'll now take the above specified model and turn it into an ECM. We shall run NW-HAC least squares model for establishing error correction mechanism.

'We now estimate the corresponding ECM:

equation table_9_4_ecm.ls(n) d(rer_def_NT) c ec1(-1) d(rer_def_NT(-1))

'Note that the SR effect is significant as the EC coefficient is of value -0.59 is statistically significant at better than 1% significance level.

''**********************************************
''S2.A & S2.B: Obtaining LR Coefficients
'***********************************************

'Now, by employing FMOLS and DOLS cointegration regression estimators, finally we shall calculate our LR coefficient i.e. BS coefficient for Korea against U.S.

equation table_9_4_LReqn1_fmols.cointreg(method=fmols) rer_def_NT rer_def_T rop gexp gdppc a_tilde

equation table_9_4_LReqn2_dols.cointreg(method=dols, trend=constant, lag=1,lead=1 ) rer_def_NT rer_def_T rop gexp gdppc a_tilde

'The BS coefficient obtained through FMOLS and DOLS estimators are statistically insignificant. Thus, there is NO evidence in support of BS effect existing for Korea.

''******************************************
'Multivariate Cointegration Approach
'******************************************

''*************************************************************
''Check if the VAR(2) model is dynamically stable
'*************************************************************
```

```
freeze(table_9_4_var2_varstable) var2.arroots(graph)

'The model is dynamically stable.

''************************************************************************
''M1.A & M1.B: Identifying the number of cointegrating vectors
'*************************************************************************
'Having identified the appropriate number of lags to put in, I now go on to test for the appropriate number of cointegrating equations.

freeze(table_9_4_var2_coint) var2.coint(s,1)

'This command estimates all possible combinations of constants and trends in the level data series and the cointegrating equations. All the results indicate 1 cointegrating vectors.
'
'GENERAL NOTE:, in practice, cases 1 and 5 are rarely used. One should use case 1 only if one knows that all series have zero mean. Case 5 may provide a good fit in-sample but will produce implausible forecasts out-of-sample. As a rough guide, use case 2 if none of the series appear to have a trend. For trending series, use case 3 if you believe all trends are stochastic; if you believe some of the series are trend stationary, use case 4.

'Note that the 5 cases are identified under "Johansen cointegration test" in EViews. They run from most restrictive (no constants in either the level series or CEs) to most general (trend terms in both the level series and CEs).

''********************************************************************
''M2.A, M2.B & M3: Vector Error Correction Model (VECM)
'*********************************************************************

' For estimating the LR relationship, corresponding VEC command is:

var table_9_4_vec_c.ec(c,1)  0 0 rer_def_NT rer_def_T rop gexp gdppc a_tilde
var table_9_4_vec_d.ec(d,1)  0 0 rer_def_NT rer_def_T rop gexp gdppc a_tilde

'CONCLUSION:  I conclude that rer_def_NT and a_tilde are not cointegrated in the Korea's data.
```

CHAPTER 10: CONCLUSION

The thesis undertakes a study of three inter-related dimensions of productivity-real exchange rate linkage under the theoretical framework of Balassa-Samuelsson hypothesis for a group of ten ASEAN and SAARC economies: (a) identification of productivity as a key determinant of permanent deviations in long-run real exchange rate, (b) validity of hypothesis under alternative theoretical specifications, and (c) inclusion of demand-side shocks to see if this can bring any substantial improvements to earlier estimates.

I conduct a careful examination to verify the long-run association between sectoral productivity imbalances and long-run real exchange rate deviations as predicted by the Balassa-Samuelson model. To assess the robustness of my results, I take into account the distinction between traded and non-traded sectors of the real economy in a more definitive manner. Data inconsistencies across sectors as well as across countries in the form of uncommon data sources, inconsistent scheme of sectoral division and inadequately disaggregated sectors are addressed to ensure data reliability. Furthermore, I employ two alternative schemes of sectoral classification to examine the sensitivity of model estimates. I also use three alternative measures of real exchange rate that dominantly comprise of non-tradable prices so that the internal mechanism of the Balassa-Samuelson model could be captured appropriately. I test three alternative theoretical specifications of the Balassa-Samuelson model, ranging from the most restrictive domestic version of the model to the modified version allowing for deviations in tradables prices from long-run PPP. Finally, I apply various time-series and panel data econometric techniques, ranging from single equation to multivariate cointegration approaches, to test the consistency and robustness of my model estimates.

The results suggest that inter-country divergent sectoral productivity patterns do not exert any significant effect on the long run real exchange rates for the ASEAN and SAARC countries in my sample as predicted by Balassa (1964) and Samuelson (1964). The argument of the Balassa-Samuelson hypothesis that biased relative productivity of tradables at home will influence the overall price level of the country through non-traded sector prices and contribute to the long-run movements of real exchange rates does not hold. My findings are highly robust against alternative sectoral classifications, different variants of real exchange rate measures, alternative theoretical specifications of the model and different econometric techniques. Empirical results for the standard (international) version of the hypothesis reveal that relative sectoral productivity differences across countries are inadequate in explaining the trend departures in the real exchange rates away from their long run equilibrium.

The Balassa-Samuelson hypothesis is more convincingly rejected when the domestic version of the model is tested. There are absolutely no traces of statistical significance in support of the basic building block of the hypothesis. I did not obtain any statistical support in favour of the notion that domestic sectoral price movements are driven by divergent productivity patterns in the long run. The country-by-country analysis, as well as panel data estimations, reveal that relative sectoral prices and productivities are not cointegrated in the domestic version of the model. This finding is also consistent with my finding for the standard (international) version of the Balassa-Samuelson hypothesis.

These results hold true when I relax the assumption of PPP in the tradable sector for the Balassa-Samuelson hypothesis. In the modified version of the model, I allow for trend deviations in international tradable prices from its PPP value. I did not find any valid existence of the Balassa-

Samuelson effect for the countries in my sample. Furthermore, the empirical results reveal a clear departure of tradable prices from the long run PPP suggesting that this divergence is a potential reason for the non-existence of the Balassa-Samuelson effect. The long run coefficient of inter-country tradable price gap capturing deviations from PPP is found to be significantly different from zero in almost all the cases. This invalidates the fundamental assumption of tradable sector PPP in the Balassa-Samuelson model.

I then attempt to model the productivity-real exchange rate association in a more inclusive theoretical framework by including a few demand-side factors that drive real exchange rate movements. The two demand-side factors that are widely recognized in the literature for inducing sizeable misalignments in the real exchange rate are GDP per capita and government consumption spending. However, their representation in the Balassa-Samuelson framework does not bring any significant difference to my previous result on productivity-real exchange rate relationship. The Balassa-Samuelson hypothesis is invalidated for a large number of individual countries. There is some support when the model is tested through a panel multivariate cointegration procedure. However, the results are not robust as the two panel cointegration regression estimators, panel FMOLS and DOLS, yield contradictory results.

On the whole, I tend to reject the Balassa-Samuelson effect for emerging Asian countries due to inadequate empirical evidence in support of the hypothesis. Irrespective of alternative sectoral divisions, real exchange rate measures, model specifications or estimation procedures, sectoral productivity patterns are rarely found to cause real exchange rate appreciation. This calls for exploring the productivity-real exchange rate relationship in future research under more realistic theoretical settings that are compatible with modern approaches to open economy macroeconomic behaviour. As a future line of inquiry, one should look into the New Open Economy Macroeconomic (NOEM) models for this purpose, and analyse short term and long term real exchange rate behaviours under imperfect goods and factor market conditions. The model should take into account idiosyncratic consumer preferences, heterogeneous composition of consumption basket across countries and nominal labour market rigidities to understand the real exchange rate and productivity relationships. However, this was beyond the scope of my dissertation.

My study suffers from a few limitations and provides room for future research. First, I fail to include low income East and South Asian economies due to non-availability of data. This might have reduced the generalization of my findings for low income economies in Asia. Second, the unavailability of data on capital formation and other potentially important factor inputs prevented me from employing more advanced measures of productivity like total and multifactor productivities. Finally, longer time-series data for the proposed models could have resulted in somewhat different time-series estimates.

For those, who intend to replicate my results or wish to conduct future explorations on the subject, my original data sets (Microsoft Excel spreadsheets), STATA program codes for sectoral data transformations and EViews program codes for empirical estimations are publicly available at Harvard Dataverse research repository. Users may access the data and programming codes using the following DOI doi:10.7910/DVN/2RMBET.

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